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Testing Volatility Asymmetry in Istanbul Stock Exchange

Cem Payaslıoğlu

Monetary Transmission and Bank Lending In Turkey

Lokman Gündüz

Deseasonalizing Macroeconomic Data: A Caveat to Applied Researchers in Turkey

C. Emre Alper & S. Borağan Aruoba

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CONTENTS

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TESTING VOLATILITY ASYMMETRY IN ISTANBUL STOCK EXCHANGE

Cem PAYASLIOĞLU*

Abstract
In this paper three different models of daily stock return volatility in the Istanbul Stock Exchange (ISE) are estimated and compared. The mean model is represented by stock return variable predicted by a MA (1) term, the day-of-the week (Monday) dummy and the risk term which is the time-varying conditional variance with three alternative specifications: These are standard GARCH-M (1,1), EGARCH-M (1,1) and TGARCH-M (1,1) models. The latter two incorporate leverage effect into the model. Choice of the appropriate volatility model is determined by inspecting level and squares of the standardized residuals. In addition to the traditional model selection criteria diagnostic tests of Engle and Ng (1993) paper are also utilized. Estimation results revealed that 1) the asymmetry component in the leverage models are not significant. 2) Portmanteau statistics did not discriminate among the models. 3) All models passed the diagnostic tests successfully. These findings point to the necessity of further research with special consideration of other garch extensions in particular: t-distributed versions as well as non-parametric alternatives need to be studied.

I. Introduction

Study of volatility in financial and economic time series has been an active field of research up to date. There are several reasons that one may need to be concerned with volatility. First, typical investor may need to deal with volatility to understand the risk of holding an asset or the value of an option. Second, forecast confidence intervals may be time varying so that more accurate intervals can be obtained by modeling the variance of the errors. Third, more efficient estimators can be obtained if heteroskedasticity in the errors is handled appropriately.

While most researchers agree that volatility is predictable, they differ on how this volatility predictability should be modeled. In recent years the

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evidence for predictability has led to a variety of approaches, some of which are theoretically motivated, while others are simple empirical suggestions. The most interesting of these approaches are the ‘asymmetric’ or ‘leverage’ volatility models, in which good news and bad news have different predictability for future volatility. (Pagan and Schwert, 1990) provide comparison of volatility models. In this paper three different models are utilized for estimating volatility in stocks traded in the Istanbul Stock Exchange (ISE). Studies pertaining to the ISE have so far found substantial evidence of volatility (Balaban, 1999). There have even been attempts to explain this volatility by macroeconomic fundamentals although with no success (Gunes et al., 1998). The rest of the paper is organized as follows: Next section introduces autoregressive conditional heteroscedasticity class of models and two asymmetry-oriented extensions. Data and methodology with respect to forming the mean model is then explained. Conventional and sign - size bias tests are exposed in the diagnostic tests section. Empirical results and conclusion then follows.

II. Arch Models and Leverage Effects

To capture the effect of the changing volatility in a time series, Engle (1982) developed the ARCH model where the conditional variance is linear function of the past squared errors, as well as possible exogenous variables X. Several extensions of the model have been subsequently developed. The ARCH models are designed to model and forecast the variance of a dependent variable. In each case the variance of the dependent variable is specified to depend upon past values of the dependent variable using some formula and or upon some exogenous or independent variables. The general representation of the model is ARCH (p) with a form:

$$h_t = \varpi + \sum \alpha_i \, \varepsilon^2_{t-i} \tag{1}$$

The ARCH model is based on an autoregressive representation of the conditional variance. One may also in an unusual way add a moving average part. The GARCH process generalized autoregressive conditionally heteroscedastic are thus obtained Bollerslev (1986). The model is defined by:

$$h_t = \varpi + \sum \alpha_i \, \varepsilon^2_{t-i} + \sum \beta_j \, h_{t-j} \tag{2}$$

Despite the apparent success of these models, they can not capture some important features of the data generating process. The most interesting of these is the leverage or asymmetric effect discovered by Black (1976). He and some other researchers have found evidence that stock returns are negatively correlated with changes in returns volatility, i.e. volatility tends to rise in response to ‘bad news’ (excess returns lower than expected) and to fall in response to ‘good news’ (excess returns higher than expected). GARCH models, however, assume that only the magnitude and not the positivity or the negativity of unanticipated excess returns determines feature of ht. The EGARCH model proposed by Nelson (1991) is an alternative which accommodates the asymmetric relation between stock returns and volatility. Moreover, the conditional variance is expressed in terms of logarithms becomes a natural device for ensuring that ht remains nonnegative. The specification for the variance is:

$$\log\left(h_t \right) = \varpi + \beta \log\left(h_{t-1} \right) + \alpha \left[\left| \frac{\varepsilon_{t-1}}{\sqrt{h_{t-1}}} \right| - \sqrt{\frac{2}{\pi}} \right] + \gamma \frac{\varepsilon_{t-1}}{\sqrt{h_{t-1}}} \tag{3}$$

Another model allowing for leverage effect is TARCH or Threshold ARCH introduced by Zakoian (1994) and Glosten et al., (1989) independently. The model for the variance is:

$$h_t = \varpi + \alpha\varepsilon^2_{t-1} + \gamma\varepsilon^2_{t-1}\, d_{t-1} + \beta\, h_{t-1} \tag{4}$$

where d_{t-1} is a dummy being equal to1 if $\varepsilon_{t-1} < 0$ and zero otherwise.

There are several approaches to testing the adequacy of an ARCH model. In some respects, the easiest is to reestimate the model with a more complex model and see whether the new parameters are significantly different from zero. Comparing each model’s log-likelihood values can be another criteria for determining appropriate model. However there are some formal diagnostic tools which can be applied to ARCH models.

III. Diagnostic Tests

Several of these are formulated in terms of standardized residuals, which are the conventional residuals divided by their one step ahead conditional standard deviation. If the model is correctly specified, these should be independent identically distributed mean zero and variance one series. They might also be normally distributed although this is not essential.

The correlogram calculated for the standardized residuals is a test for remaining serial correlation in the levels and checks whether the mean is correctly specified. Similarly, one can use the histogram of standardized residuals. Here the kurtosis is now 6 when it was 30 for the original series. A test for asymmetry can be based on the Ljung Box test for skew (cross) correlation. Simply one may look at the cross correlogram between the level and the square of the series to determine whether there is predictability of the square by the level. This can help decide when a leverage effect might be needed. A successful model should not have a significant Ljung-Box statistic for the skew correlations in the standardized residuals. Cross correlogram should be zero for GARCH and negative for EGARCH and TGARCH type models.

The models are estimated by maximizing the log-likelihood function, assuming that is conditionally normally distributed. Even if this assumption is not correct, as long as the conditional mean and variances are correctly specified, the quasi maximum likelihood estimates are consistent and asymptotically normal, as pointed out by Glosten et al. (1989) and (Bollerslev and Wooldridge, 1992). Detection of excess skewness and kurtosis in the residuals reveal whether the models are correctly specified. The null hypothesis is that errors are drawn from a conditional normal distribution. Campbell and Henschell (1992) have previously applied these tests to GARCH-M models. Another diagnostic is to examine the standardized residuals from the estimated models. They should be independently and identically distributed. Engle and Ng (1993) developed diagnostic tests known as Sign Bias Test, Negative Size Bias Test and Positive Size Bias Tests. These could be indicative of misspecification.

In the Sign Bias Test, the squared standardized residuals are regressed on a constant and a dummy variable, denoted S_{t-1}. This is an indicator variable that takes a value of one if ε_{t-1} is negative and zero otherwise. Sign Bias Test statistic is the t-statistic for the coefficient of S_{t-1}. This test shows whether positive and negative innovations affect future volatility differently from the prediction of the model. In the Negative Size Bias Test the squared standardized residuals are regressed on a constant and the variable denoted as $S_{t-1}\varepsilon_{t-1}$. The t-statistic for the coefficient of this variable is indicative of whether larger negative innovations are correlated with larger biases in future volatility. In the Positive Size Bias Test the squared standardized residuals are regressed on a constant and the variable denoted as $S^{+} \varepsilon_{t-1}$ where $S^{+} = 1- S_{t-1}$. Significant t-statistic shows that larger positive innovations are correlated with larger biases in future volatility.

IV. Data and Methodology

The data consists of daily stock returns based on closing prices of index (ISE-100) comprised of 100 most heavily traded stocks during the period between January 1, 1990 and July 20, 2000. Originally data starts on January 1, 1986, however, because the volume of trade was too small in the early days and trading occurred only on one day of the week (Fridays) there was no substantial volatility. Therefore beginning date is changed to January 1, 1990. Consequently the sample is reduced from 3255 observations down to 2611. The existence of a probable ARCH (GARCH) effect, can best be indicated by the correlogram based on the squared stock return series.

4.1. Moving Average Representation of the Mean Equation

Researchers studying stock market behavior have generally found some patterns such as nonsynchronous trading in individual stocks, bid-ask spreads, and minimum size price changes can cause serial correlation in stock and index returns. Since these can induce a small, short-lived serial correlation in these returns, while the ARCH models assume that conditional error is serially uncorrelated, it is necessary to extract this serial correlation from the stock return mean (Hamao et al., 1990). This means that conditional mean return takes a moving average component so the equation becomes:

$$R_t = a + b\ \varepsilon_{t-1} + \varepsilon_t \tag{5}$$

4.2. Day-of-the Week Effect

Daily stock returns generally exhibit high variability and studies made for various stock markets found higher stock returns for Mondays. To take this potential Monday effect into consideration, a dummy variable representing Monday is added to the mean equation. (Usually the dummy is inserted into both mean and variance equations, however it did not turn out to be significant in the variance equation). Thus the model becomes:

$$R_t = a + cD_t + b\varepsilon_{t-1} + \varepsilon_t \tag{6}$$

4.3. Measuring Effect of the Risk on Return

A particularly interesting predetermined variable to introduce in this model is the conditional variance (or the conditional standard deviation) itself. This is an attractive form in financial applications since it is natural to suppose that the expected return on an asset is proportional to the expected risk of the asset. Measuring this by the conditional variance leads to the ARCH-M (ARCH in mean) model. This is an extension of GARCH model to allow the conditional mean to be a function of conditional variance at time t. Finally the mean equation becomes:

$$R_t = a + dh_t + cD_t + b\varepsilon_{t-1} + \varepsilon_t \quad (7)$$

V. Empirical Results

Table 1: Summary Statistics for the ISE-100 Series

Number of observations	2611
Mean	0.00297
Median	0.00201
Maximum	0.35604
Minimum	-0.1794
Standard deviations	0.03239
Skewness	0.46439
Kurtosis	10.3282
Jarque-Bera statistic	5936.18

Summary statistics reveal important characteristics. The sample mean return is positive and a significant departure from normality can be observed. Although the mean and median values are close, the series are revealed to have positive skewness and leptokurticity. This is to be expected as a typical characteristic of many financial series. In addition, Table 2 reveals that Ljung-Box statistic based on the correlogram of the levels and squares of the stock returns are highly significant. This statistically confirms the existence of both serial correlation and time-varying variance. It appears that autocorrelation has been successfully removed by short-lived serial correlation term, that is MA(1) in the mean regression part. However, unpredictable part still pointed to probable existence of heteroscedasticity as confirmed by squared residuals exhibiting a significant Q-statistic. Thus, ARCH (GARCH) type of modeling to capture time varying variance is called for.

Table 2: Ljung-Box Statistics for the Original Series and Residuals

Ljung-Box Q statistic	Levels	Squares
Stock return series	42.748*(0.00)	236.95*(0.00)
Mean without GARCH term	16.114 (0.137)	196.15*(0.00)
Mean with GARCH term	17.682(0.089)	11.625(0.392)}

Note: Values in the second row based on residuals from the mean regression model excluding garch term and those on the third row based on standardized residuals from the model with GARCH (1,1) term. *Significant at 1 % level.

Table 3: Estimation Results for GARCH-M (1,1)

Mean Model
$R_t = 0.00004 + 2.61539h_t + 0.00240D_t + 0.12766\varepsilon_{t\text{-}1} + \varepsilon_t$ (0.00087) (0.93982) (0.00115) (0.02257)
Conditional Variance
$h_t = 0.00005 + 0.17356\ \varepsilon^2_{t\text{-}1} + 0.77828\ h_{t\text{-}1}$ (0.00109) (0.02665) (0.02849)   Logl=5545.09

Note: Parameters are estimated using maximum likelihood with Marquard algorithm. Bollerslev-Wooldridge robust standard errors are given in parenthesis.

In the mean regression equation all parameters except the constant term are significant thus statistical evidence of day-of-the week as well as risk effect can be detected. In the variance equation on the other hand, both arch and garch terms are found significant. Both terms are positive thus fulfilling the positivity requirement. On the other hand, the sum of estimated α and β, although not equal to unity, is around .95. This implies that the conditional variance is quite persistent. Following estimation, standardized residuals were generated. If the model is correctly specified, these should be independent identically distributed mean zero and variance one series. They might also be normally distributed although this is not essential.

Table 4: Estimation Results for EGARCH-M (1,1)

Mean Model
$R_t = \underset{(0.00090)}{0.00063} + \underset{(0.98592)}{3.46735h_t} + \underset{(0.00125)}{0.00301D_t} + \underset{(0.02213)}{0.13615\varepsilon_{t-1}} + \varepsilon_t$
Conditional Variance
$\log(h_t) = \underset{(0.16993)}{-0.82842} + \underset{(0.01877)}{0.92106\log(h_{t-1})} + \underset{(0.02849)}{0.35158}\left[\left\lvert\frac{\varepsilon_{t-1}}{\sqrt{h_{t-1}}}\right\rvert - \sqrt{\frac{2}{\pi}}\right] \underset{(0.02914)}{- 0.02228}\frac{\varepsilon_{t-1}}{\sqrt{h_{t-1}}}$
Logl=5549.61

Estimation with EGARCHM, like previous model, produces a good mean equation. On the other hand, the parameter representing the asymmetry term in the variance i.e., $\varepsilon_{t-1} / \sqrt{h_{t-1}}$ is not significant.

Table 5: Estimation Results for TGARCH-M (1,1)

Mean Model
$R_t = \underset{(0.000897)}{0.000054} + \underset{(0.936910)}{2.516648h_t} + \underset{(0.001166)}{0.002467D_t} + \underset{(0.022395)}{0.129996\varepsilon_{t-1}} + \varepsilon_t$
Conditional Variance
$h_t = \underset{(0.00001)}{0.000056} + \underset{(0.027961)}{0.160153\varepsilon^2_{t-1}} + \underset{(0.041288)}{0.026359\varepsilon^2_{t-1}d_{t-1}} + \underset{(0.028679)}{0.776664h_{t-1}}$
Logl=5545.65

The same pattern as in the previous case can be observed regarding the mean equation. However, in the variance equation the coefficient of the term representing leverage i.e., $\varepsilon^2_{t-1}d_{t-1}$ is not significant. This casts doubt on the adequacy of the model above. Among these models EGARCH-M has the highest log-likelihood value. Therefore standardized residuals-based diagnostic tests is called for.

Table 6: Portmanteau and Sign Type Diagnostic Statistics for the Models

Model	LB(12)	$LB^2(12)$	Sign	Negative	Positive
GARCH-M(1,1)	18.259	24.523	1.686	-1.085	-0.481
EGARCH-M(1,1)	20.996	22.669	1.395	-0.453	0.049
TGARCH-M(1,1)	19.022	22.896	1.424	-0.721	-0.196

Notes: This table reports the diagnostic test results of three alternative volatility models for the daily return of the ISE-100 Index. The first and second columns refer to Ljung-Box portmanteau statistics for the twelfth order serial correlations in levels and squared standardized residuals, respectively. Since the mean regression model has one MA (1) term, statistics are compared against X^2 (11), that is twelve minus one degrees of freedom.

Comparison of the models on the basis of correlogram of standardized residuals did not produce a strong argument in favor of any of these models. One thing to note is that egarch-m standardized residual was significant at 5 % although not at 1% level. The controversy is more pronounced when the squares of the standardized residuals are inspected. In this case the asymmetric models look better as they have insignificant statistics both at 1 and 5% level.

All the models pass the sign bias, negative size bias and positive size bias tests successfully. In calculating these tests standardized residuals from the three models are squared and regressed on the sign and size variables. These are calculated on the basis of ε_{t-1} obtained from each model as indicated above.

Finally, cross-correlogram between standardized residuals and their squares for each model do not produce significant Q statistic. The results, calculated for both lead and lag, are shown in Table 7.

Table 7: Cross-Correlogram Between Standardized Residuals and Squares

Model	Lag	Lead
GARCH-M	10,489	15,271
EGARCH-M	9,073	14,994
TGARCH-M	10,419	15,173

Correlations are asymptotically consistent approximations. Lag correlation is correlation between square of standardized residual at time t and standardized residual in level terms at time t-i, with i being the lag (lead). The latter becomes t+i for the lead. These are calculated for every i =0 through 12.

VI. Conclusion

Estimation results statistically confirm the existence of day-of week effect and short lived serial correlation in the innovations. This seems compatible with similar findings for many other markets. Risk and return are positively related for all models and the relationship is documented as statistically significant. On the other hand, empirical results do not support any prevailing asymmetry. The evidence can be found in the following: 1) alternatives to GARCH-M allowing for leverage do not outperform it on the basis of standard diagnostic tests. 2) The asymmetry term in alternative models are not significant 3) Sign and size-based specification test statistics using standardized residuals do not favor any one of these models as the best. In addition the log-likelihood of the all models are very close although that of EGARCH-M is slightly higher. This might lead the conservative researcher to retain GARCH-M . However, developments of GARCH extensions are numerous and only two popular alternatives are investigated here. Besides, on account of the leptokurticity of the series, econometric modeling under normal distribution assumption generally do not provide solid groundwork. Consequently, there may be a need of using other alternatives under different distributional assumptions as well as non-parametric models in order to have an exhaustive coverage of the subject.

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MONETARY TRANSMISSION AND BANK LENDING IN TURKEY

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Abstract

This paper analyses the role of bank lending in the monetary transmission mechanism in Turkey. We present evidence from a VAR-model estimated with monthly aggregate data covering the period 1986-1998. We find some evidence for the existence of the bank lending channel, though inadequate given the identification problem. We observe that following a monetary contraction, aggregate bank credit and securities holdings of the banks decline immediately much more than the money (deposits) does. The timing of impulse responses of the credits and output, and the results of variance decomposition seem to favour bank lending view. Moreover our results are also consistent with the traditional interest rate channel and the exchange rate channel.

I. Introduction

In recent years, a large of literature has developed that emphasises the role of credit market imperfections in the monetary transmission process, known as the "credit view". Part of this literature focuses on the existence and importance of a bank lending channel. The implications of Turkish institutional setting for the effectiveness of monetary policy through bank lending are ambiguous. On the one hand, relations between banks and firms, which lie at the core of credit view, are relatively strong in Turkey. Turkish banking system is the main source of (although not the only one) external finance especially for short-term maturities. Importance of reserve requirements as a direct control mechanism also provides another

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motivation for the investigation. On the other hand, because of financial liberalization process, the importance of the traditional bank lending channel would diminish over time. Especially financial deregulation and innovation could lead to important changes in the financial structure. Even the effectiveness of monetary policy for the case of Turkey may well be questioned, given the chronic problems of Turkish economy including high inflation rates, currency substitution and huge budget deficits etc.

Although the importance of financial market imperfections in the monetary transmission process (as predicted by the credit view) has been established by a large number of studies, the empirical evidence for the existence of a bank lending channel has been much less conclusive. To a large extent, this is due to the fact that most studies based on aggregate data suffer from a severe identification problem. This is the inability to establish whether the decrease in credit that is observed after a monetary contraction is induced by bank supply or driven by a fall in borrowers' demand. In the latter case, a lending channel would be irrelevant. In this respect, recent studies based on disaggregate data show that the responses of credit variables can be analysed in combination with other hypotheses that follow from the theoretical literature underlying the credit view. Information problems, for instance, are presumably more relevant for particular categories of borrowers.[1] It appears from this line of research that following a monetary contraction, the amount of bank credit to small firms is reduced while large firms initially attract more (mostly short-term) credit as a buffer to compensate for declining cash flows. Yet, although this is obviously consistent with the credit view in the sense that credit is `special,' there still is no general agreement to what extent these findings should be interpreted as self-evident support for a bank lending channel.[2]

Unfortunately, detailed time series at the individual firm or bank level are not available for most countries. Empirical research may still yield valuable insights, though, even if based on aggregate data.[3] For example, following Bernanke and Blinder [1992], Yulek [1998] shows the importance of bank lending channel for Turkey.

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1 See Gertler and Gilchrist [1993, 1994]; Oliner and Rudebush [1996].

2 See Oliner and Rudebush [1996] and Kashyap et al. [1996] for a discussion.

3 Studies at a sectoral level may be useful, albeit less rigorous, alternative. In most of the studies, bank lending is split into loans to corporate sector and loans to the household sector See e.g. Dale and Haldane (1995) for the United Kingdom, Kakes (1998) for the Netherlands, and Kakes, Sturm and Maier (1999) for Germany.

The purpose of this paper is to provide more evidence on the role of banks in transmission of monetary policy in Turkey, using vector autoregression (VAR) analysis over the period 1986-1998. Our analysis is an extended study of Ekren and Gündüz (1999) and similar to Garretsen and Swank (1998), who also use aggregate data in their study for the Netherlands. Since the identification problem still persists, we find some evidence for the existence of bank lending view in Turkey. Moreover our results are also consistent with the traditional interest rate channel and the exchange rate channel.

The remainder of the paper is organised as follows. In Section II we briefly review the credit channel of monetary policy. The specification of our model and the selection and pre-testing of the data are dealt with in Section III. Section IV presents the main results and Section V concludes.

II. The Credit View of Monetary Transmission

The monetary transmission mechanism is commonly divided into the money view and credit view of monetary transmission.[4] The money view can be referred to as the traditional view of monetary transmission to real activity as described in its simplest version in standard IS-LM models. In the IS-LM model, only two financial assets are distinguished, money and bonds, of which the latter is supposed to be representative for the whole capital market. Since banks do not play an essential role in this world, there is no need to distinguish bank loans from other bank assets. According to this approach, monetary policy works primarily through its impact on the capital market interest rate.

The credit view can be termed as the capital market imperfections approach.[5] The credit view is based on the assumption that the same informational and agency problems that explain many aspects of financial structure also play a role in monetary transmission. The “credit channel” theory of monetary policy transmission holds that informational frictions in credit markets worsen during tight-money periods. The resulting increase in the external finance premium (the difference between internal and external funds) enhances the effects of monetary policy on the real

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4 Worthwhile to mention is that asset price channels have been ignored in the money versus credit debate. Mishkin [1996] recently defines five different monetary transmission channels.

5 See Bernanke and Gertler [1995] and Hubbard [1994] for an evaluation of the credit channel

economy. According to credit channel theory, the direct effects of monetary policy on interest rates are amplified by endogenous changes in the external finance premium. The size of external finance premium reflects imperfections in the credit markets that drive a wedge between the expected return received by lenders and the costs faced by potential borrowers. In the credit channel, a change in monetary policy that raises or lowers open-market interest rates tends to change the external finance premium in the same direction. Therefore, additional effect of policy on the cost of borrowing broadly defined -and consequently, on the real spending and real activity- is magnified.

The credit view complements the money view by focusing on two channels of monetary transmission mechanism. These are bank lending channel and balance sheet channel.[6] According to bank lending view monetary policy may also affect the external finance premium by shifting the supply of intermediated credit, particularly loans by commercial banks. Banks, which remain the dominant source of intermediated credit in most countries, specialize in overcoming informational problems and other frictions in the credit markets. If the supply of bank loans is disrupted for some reason, bank-dependent borrowers may not be literally shut off from credit, but they are virtually certain to incur costs associated with finding a new lender, establishing a credit relationship, and so on. Therefore, a reduction in the supply of bank credit, relative to other forms of credit, is likely to increase the external finance premium and reduce real activity.

The bank lending channel has two clear parts. First, the bank credit is special. There is no perfect substitute to bank loans, both on the liability side of banks' balance sheets and on the asset side of borrowers. Especially, households and small firms lack access to other forms of credit than bank loans. Second, monetary policy changes have a direct effect on money supply. Following a monetary tightening which drains deposits from the banking system, banks have to readjust their portfolio by reducing their supply of loans, given the imperfect substitutability between loans and other assets. Loan supply being reduced, banks increase their lending rate or reduce their loans. Thus a reduction in the supply of loan leads to a rise in the external finance premium for bank-dependent borrowers whose activity is reduced. As a result, credit allocated to bank-dependent borrowers may fall causing these borrowers to curtail their spending.

[6] According to Bernanke and Gertler [1995], the existence of a balance sheet channel seems fairly well established, while the bank lending channel is more controversial.

The balance sheet channel emphasizes the potential impact of monetary shocks on borrowers' financial position. The financial position of borrowers, their net worth, can be determined from their balance sheets and income accounts. The basic idea is that any shock affecting borrowers' financial position modifies the external finance premium and the overall terms of credit that borrowers face.

III. Specification and Pre-testing of the Model

3.1. Data and Selection of Variables

Our time series consist of monthly data covering the period January 1986-October 1998.[7] This implies that we have 154 observations for each variable. The main reason for choosing January 1986 as the starting date of the sample is related to the healthiness and the availability of data. This date also coincides with the radical changes in the Turkish financial system.

The selection of variables has been based on both economic and statistical criteria. Since the role of bank balance sheets in the transmission mechanism is in the primary focus, bank assets and liabilities should be included in any case. Hence, the VAR model contains equations for bank deposits, bank lending and securities holdings of banks, which may be evaluated as a minimum set of relations characterizing the behavior of banks.[8] In addition two variables, industrial production and the wholesale price index, pertaining to the real sector of the economy, are included, since they are the main target variables that reflect the eventual effects of monetary policy. Furthermore the real exchange rate is contained in the set of variables, so as to take the openness of the economy into account. The selection of policy variable is crucial in the model. Following Bernanke and Blinder (1992 and most of the subsequent VAR-based literature on monetary policy transmission, we have chosen overnight interbank interest rates as the policy variable.[9] The corresponding equation in the VAR can be interpreted as the central bank's reaction function, while

[7] All data were downloaded from Central Bank's web site: http://www.tcmb.gov.tr.

[8] We also included M2 instead of total bank deposits, but the results did not change.

[9] In a recent study by Kalkan M., A. Kıpıcı and A. T. Peker (1998), the interbank interest rate and the exchange rate basket were found as the strongest leading indicators of inflation. It seems clear that monetary aggregates denominated in TL have also left Turkish Central Bank as they did for many others all over the world.

innovations of the policy variable reflect unanticipated monetary policy shocks.

Now the complete set of variables reads as follows:[10]

• Wholesale price index (in logs)	logwpisa
• Industrial production index (real and in logs)	logipisa
• Real dollar exchange rate (in logs)	logfxsa
• Securities holdings of banks (real and in logs)	logsecsa
• Total bank loans (real and in logs)	logtcrsa
• Total bank deposits (real and in logs)	logtdepsa
• Overnight interbank interest rate	ibrsa

Besides the variables listed above, our VAR-model also includes a dummy variable for period February 1994-May 1994.

3.2. Pre-Testing and Ordering of the Variables

In order to find out whether the VAR-model should be formulated in terms of levels or the first differences of the variables, we have tested the above series for stationarity. Applying the rules stipulated by Dickey and Pantula (1987)[11], it can be deduced from Table 1 that all series, except overnight interbank interest rate, are I(1). This calls for differences all series once, except interest rate, so as to make them stationary.

Akaike's Information Criterion (AIC) has been carried out to determine the optimal number of lags. In most cases the number of lags equals three or four. Although bigger numbers (i.e. those bigger than 13) seem more appropriate for AIC, we tried to conserve degrees of freedom, and ended up with four as the number of lags.

We then applied the Johansen (1991) procedure to investigate cointegration. If there is cointegration between the I(1) variables, the relevant cointegrating residuals should be included as regressors next to the (lagged) differenced series (and, optionally, a constant and a time trend); otherwise, the system would be misspecified. The alternative, which is

[10] Indexes are normalized to 1987. All data, except interbank interest rate, are deflated by wholesale price index and then transformed in logs. All series are seasonally adjusted by X11 ARIMA method using additive approach. Real dollar exchange rate is defined as the ratio of foreign price index to the domestic price index multiplied by the nominal exchange rate.

[11] Dickey, D and S.G. Pantula (1987), "Determining the Order of Differencing in Autoregressive Processes", Journal of Business and Economic Statistics 5, pp. 451-61.

Table 1: Augmented Dickey-Fuller (ADF) Tests: Absolute t-Ratios

Dimension of series	**LOGWPISA**	**LOGIPISA**	**LOGRFXSA**	**LOGTSECSA**
Second difference	8.36^{**}	8.68^{**}	7.63^{**}	8.34^{**}
First difference	3.93^{*}	5.24^{**}	3.83^{*}	6.10^{**}
Level	1.91	2.69	2.34	1.79
Dimension of series	**LOGTCRSA**	**LOGTDEPSA**	**IBRSA**	
Second difference	6.58^{**}	9.66^{**}	8.47^{**}	
First difference	3.54^{*}	5.73^{**}	6.54^{**}	
Level	2.45	0.44	3.79^{*}	

Notes: Up to a maximum of 6 autoregressive terms, aconstant and a time trend, if significant at the 10% level, have been included in ADF regressions.:* (**) denotes that the hypothesis of stationarity is rejected at the 5% (1%) significance level, based on Mackinnon (1991) critical values.

commonly pursued, is to estimate the VAR-model in levels, after that cointegration has been established. But this, as Robertson and Wickens (1994) mentioned, would lead to a loss of efficiency.[12] Table 2 shows the results of the cointegration analysis based on Johansen (1991).[13] As it turns out, there are two cointegrating equations at 5% significance level. As it is seen, the estimated VAR-model is in fact a vector error correction model (VECM).

[12] Robertson, D. and M. Wickens (1994), "VAR modelling" in Applied Economic Forecasting Techniques, edited by Stephan Hall, Harvester Wheatsheaf, pp. 29-47.

[13] Johansen Soren (1991), "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", Econometrica 59, pp.1551-80.

Table 2: Johansen Test for Cointegration

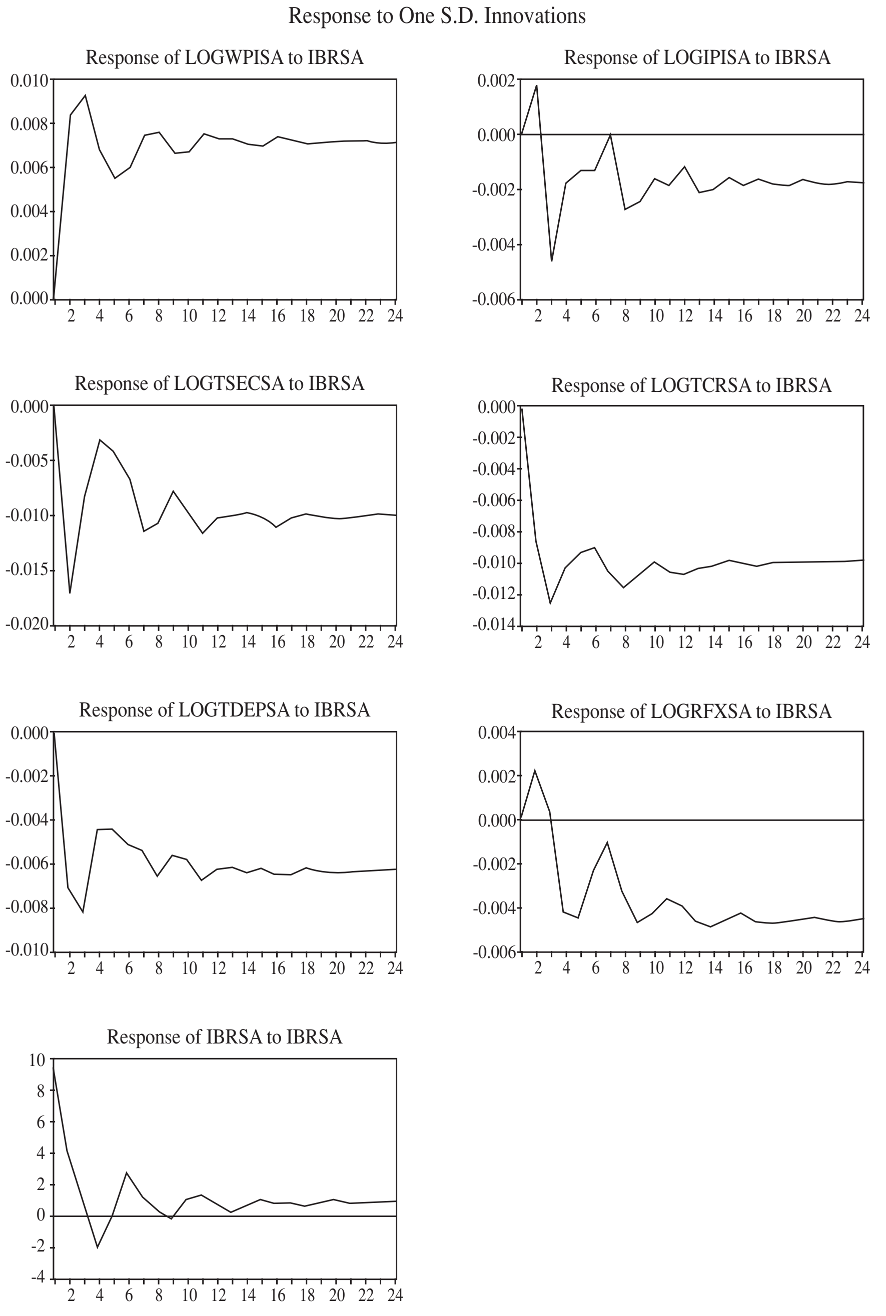
Series: LOGWPISA LOGIPISA LOGTSECSA LOGTCRSA LOGTDEPSA LOGRFXSA				
Eigenvalue	**Likelihood Ratio**	**5 Percent Critical Value**	**1 Percent Critical Value**	**Hypothesized No. of CE(s)**
0.262535	140.9528	114.90	124.75	None **
0.214781	95.57677	87.31	96.58	At most 1 *
0.147817	59.54965	62.99	70.05	At most 2
0.095944	35.71655	42.44	48.45	At most 3
0.078329	20.68777	25.32	30.45	At most 4
0.055668	8.534334	12.25	16.26	At most 5

L.R. test indicates 2 cointegrating equation(s) at 5% significance level

Notes: The test is performed under the assumption of a linear deterministic trend in the data. The underlying VAR-model has a lag length of 4 months. *(**) denotes rejection of the hypothesis at 5% (1%) significance level.: Critical values are based on Osterwald-Lenum (1992).

The primary objects of interest are impulse response functions (IRF). One commonly used method for obtaining these measures is to orthogonolize the contemporaneous innovations in the variables using Choleski factorization, which orders the innovations according to a recursive system. The order of variables indicates whether innovations in a given variable are allowed to exert a direct, contemporaneous influence on other variables in the system. Following Gertler and Gilchrist (1993), Garretsen and Swank (1998), among others, the policy variable is ordered last, implying that an innovation in the interbank interest rate has only a lagged influence on the other variables. The ordering of the equations corresponding to the list of variables reflects a transmission mechanism in which monetary policy first affects the three bank variables, then real dollar exchange rate and finally the real sector of the Turkish economy. As noted by Dale and Haldane (1995), the ordering of equations explaining the non-policy variables does not have a bearing on the effects of a shock in the policy variable as long as the latter is ordered last. In our case, as in Garretsen and Swank's (1998) study, we have seen that even ordering the policy variable before the non-policy variables, does not result in very different IRFs except the first few months of the simulation period, and hence does not lead to different conclusions.

Figure 1: Impulse Response Functions: Responses to a (One Standard Deviation) Shock in the Interbank Interest Rate



IV. Innovation Analysis and the Interpretation of the Results

Our results are based on innovations analysis (impulse response analysis and variance decompositions), which can be performed with an estimated VAR model, transformed into its moving average representation. Figure 1 depicts these IRFs generated from the estimated VECM. For each variable, the horizontal axis shows the number of months that have passed after the impulse has been given, while the vertical axis measures the response of the relevant variable, where a value of 10 corresponds to 10 % in case of interest rate, and a value of 0.002 corresponds to 0.1 % of the baseline value in case of other variables. Figure 1 thus shows that the initial shock in the interbank interest rate amounts to 10 %. The transience of the shock is evident from the fact that the interest rate increase has virtually died out after six months.

Focusing first on the effects on bank balance sheets, it turns out that the major responses to the interest rate shock are in securities holdings of the banks and total bank loans. The decrease in securities holdings is initially quite strong, running to almost 1.7 % in the second month, but at the end the of simulation period, the effect has diminished to around 1%. Bank loans have also reacted strong against the interest rate shock, declining to 1.3 % in the third month, and then keeping its level around 1 % decrease along the simulation period. Total bank deposits also decline in response to shock in interest rates, but more modest in compare to other balance sheet items. The decline reaches 0.08 % in the third month and continues at around 0.06 % in the rest of period.

The rise in interest rate causes the dollar exchange rate to appreciate modestly in the first instance but to depreciate later on.[14] The latter effect can be traced back to the sharp and persistent fall in the industrial production. Initial increase in the real activity is not statistically significant. Tight monetary policy, however, seems to work only after the second month. Induced decline in industrial production amounts to 0.05 % in the

[14] For an exchange rate channel to exist, two conditions must be fulfilled. First, the shock in the interest rate should result in an appreciation of the currency. Second, the appreciation should generate a decline in output and prices. According to simulation results, a rise in the interest rate results in an appreciation of the currency (decrease in the exchange rate) after an initial rise. An "exchange rate puzzle" is initially observed for the rise in the money market rate is followed by a depreciation, probably because the rise in the interest rate was aimed to defend the parity, but was not sufficient to counter speculative attacks. Figure 2 in the appendix shows how exchange rate generate a decline in industrial production and inflation.

third month, and tends to stay at 0.02 % at the end of simulation period.

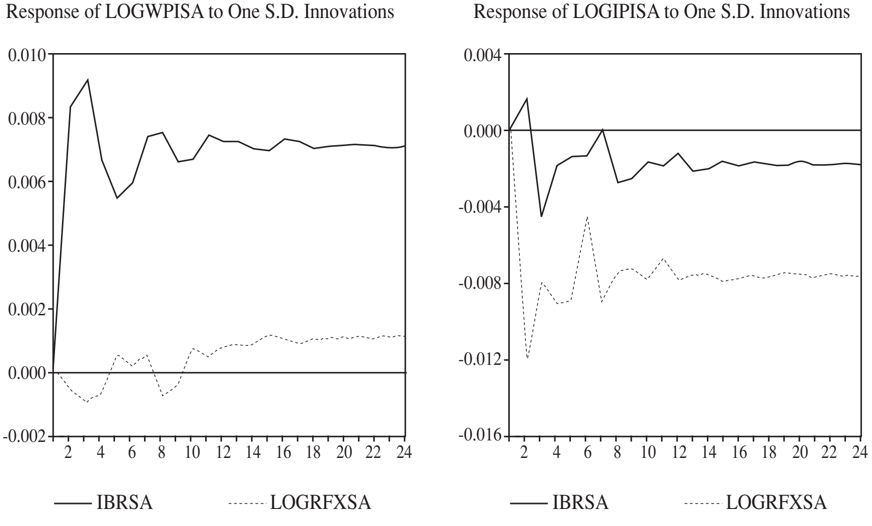
Prices show a positive response, which is in contradiction with the expected result of a monetary contraction. Kalkan, Kıpıcı and Peker (1998) also point out the unexpected positive sign between interbank interest rate and inflation, and offer some plausible explanations for Turkey. These are cost-push or wealth effect type mechanisms, the role of inflationary expectations, and the degree of effectiveness of monetary policy in a dollarized economy with inertial inflation. On the other hand, variance decompositions indicate that interbank interest rate explains much more of the variations in the wholesale price index than vice versa, hence stresses the role of cost mark-up pricing.[15]

This perverse response of prices shows up in many VAR-based studies. Sims (1992) suggests including oil price or a commodity price index in order to take supply effects into account. Dale and Haldane (1995) suggest that the positive response of prices after a monetary tightening may be explained by increasing variable costs, which initially translate into higher prices due to cost mark-up pricing.

The question arises as to how our findings fit in the discussion on money versus credit view. One approach to identify a bank lending channel is to see how banks alter their assets and liabilities during periods of monetary restraint. Accordingly, a number of studies have examined how banks adjust loans, securities, and deposit and non-deposit liabilities to changes in monetary policy. Mostly following Bernanke and Blinder (1992) and Romer Romer (1990), several stylized facts about bank portfolio behavior emerged from this line of research. First, in response to a tightening of policy, bank transactions deposits or core deposits fall immediately. Second, total bank loans decline, but only after a significant lag of two to three quarters. Third, banks are able to maintain lending in the face of a decline in core deposits by selling securities and by issuing managed liabilities such as time deposits and Eurodollar borrowings. Fourth, the eventual decline in bank lending is roughly contemporaneous with a decline in economic activity as measured by industrial production or GDP. Taken as a whole, the time lags in bank lending declines and contemporaneous decline in loans which is also consistent with a decline in output causing a fall in loan demand, pose serious questions for the existence of a credit channel.

[15] In our case, results of Granger causality tests show that relationships between interbank interest rate and inflation run in both directions. See Gündüz (1999).

Figure 2: Impulse Response Functions: Responses to a (One Standard Deviation) Shock in the Interbank Interest Rate and Exchange Rate



According to the money view, a tightening of monetary policy leads to a fall in the money supply and a subsequent contraction of nominal income. The results of VAR-analysis are fully in line with these conjectural effects. As bonds and bank loans are considered perfect substitutes in the money view, these assets are expected to respond identically to an interest rate shock. This is, to a certain extent, contradicted by our findings. Initially banks reduce both securities and loans immediately. The decline in securities is especially more dramatic. But then the decline in securities holdings of the banks take an opposite direction, and the decline diminishes to a low level, while the decline in bank loans remains high.

It is however contentious to draw a conclusion from this result in order to support the credit view. The stylized fact, that the securities holdings of the banks initially reduce much faster than the loans implying sort of a buffer-stock behavior for initial portfolio adjustment, is not very well documented in our example.[16] Instead, Figure 1 shows clearly that banks

[16] Bernanke and Blinder (1992) have pointed out that the fact that loans do not immediately fall as a reaction to monetary contraction is in itself no evidence against the credit view. If loans are quasi-contractual arrangements that are hard to change in the short run, the necessary initial portfolio adjustment is instead undertaken by shedding the more liquid bonds.

reduce their credits as a reaction to a policy tightening. In other words, there are no long significant time lags in bank loans after monetary contraction.

In the empirical testing of the bank lending view, the central problem is to identify whether the movement of credit can be explained by the demand side or by the supply side. The conclusion that the timing of the responses of loans and output to monetary policy innovation is similar is met with criticism in the literature as a source of evidence. After all the response of bank loans might just as well, have been triggered by the traditional interest rate channel. However in a recent study based on a VAR-model, Yülek (1998) come to conclusion that there is a bi-directional relation between output and credit and that a monetary shock has a relatively important effect on credit. In our study, it seems difficult to establish such a bi-directional relationships from the results of impulse response functions, although there seems to be sort of similarity regarding timing of responses, and even if the decline in bank loans takes place before the output decline. In that respect variance decompositions could be more informative. Figure 3 shows that near 30 % of the variations in the industrial production is explained by total credits, while only about 6 % of variations in the credits is explained by the output.[17]

On the other hand, the information content of credit about developments in the real economy is expected to loose its importance given highly dollarized economy. Figure 4 shows the coefficients of correlation between the impulse responses of the three financial variables on the one hand and the impulse responses of inflation and industrial production with leads to 1 to 6 months on the other hand. As it turns out, credits performs relatively well for the information content about future industrial production and inflation in comparison to other financial variables. In other words, bank loans relatively speaking include better information for future economic activity.

It should be noted that financial liberalization, institutional changes, as well as high inflation led changes in the definition of money in Turkey. Standard monetary aggregates are expected to lose their importance in an environment where currency substitution is extremely widespread. Therefore this caution should be also reserved for the bank credits as well.

[17] Bivariate and multivariate tests of Granger causality also reveal the direction running from bank credits to output. Certainly this does not imply the causality in the vocabulary form.

Figure 3: Variance Decompositions

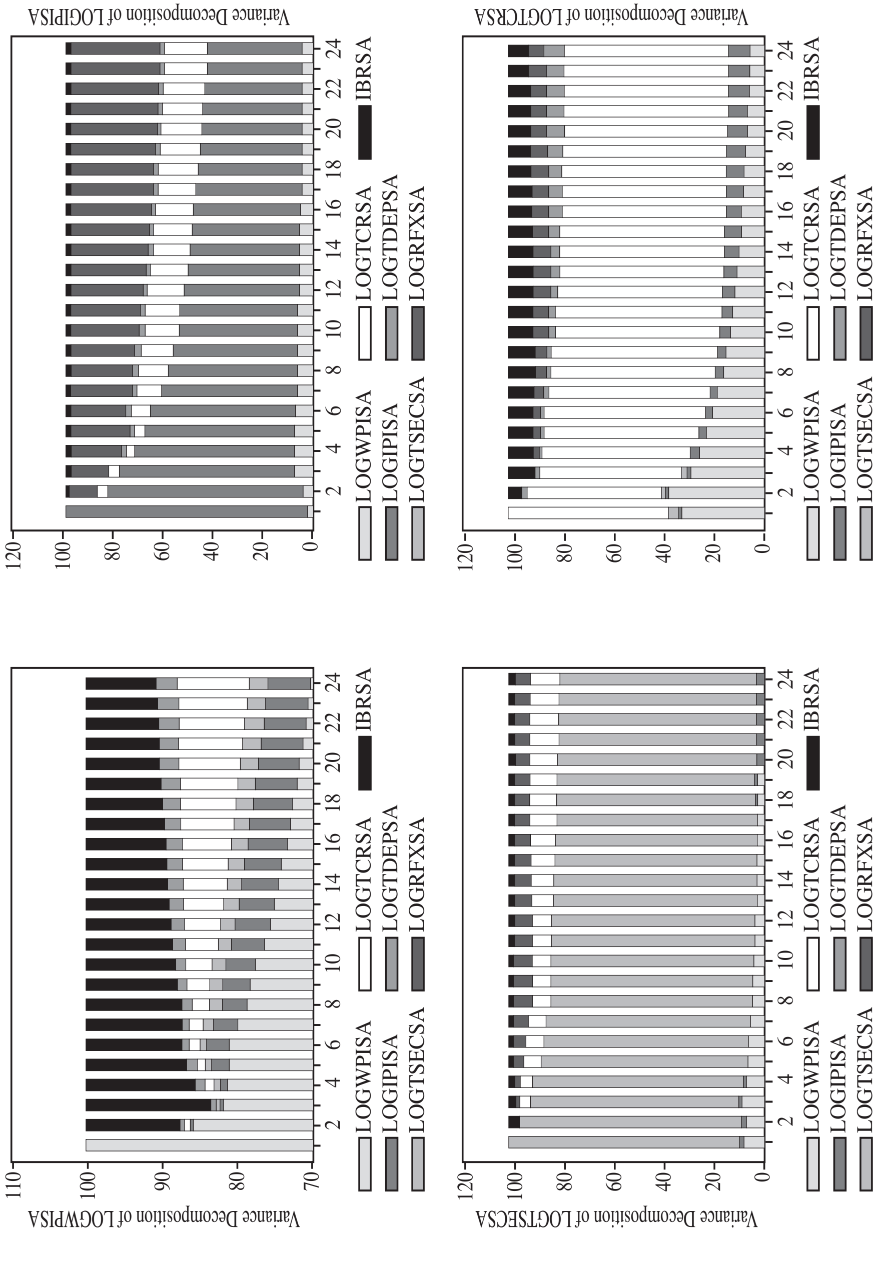


Figure 3: Variance Decompositions

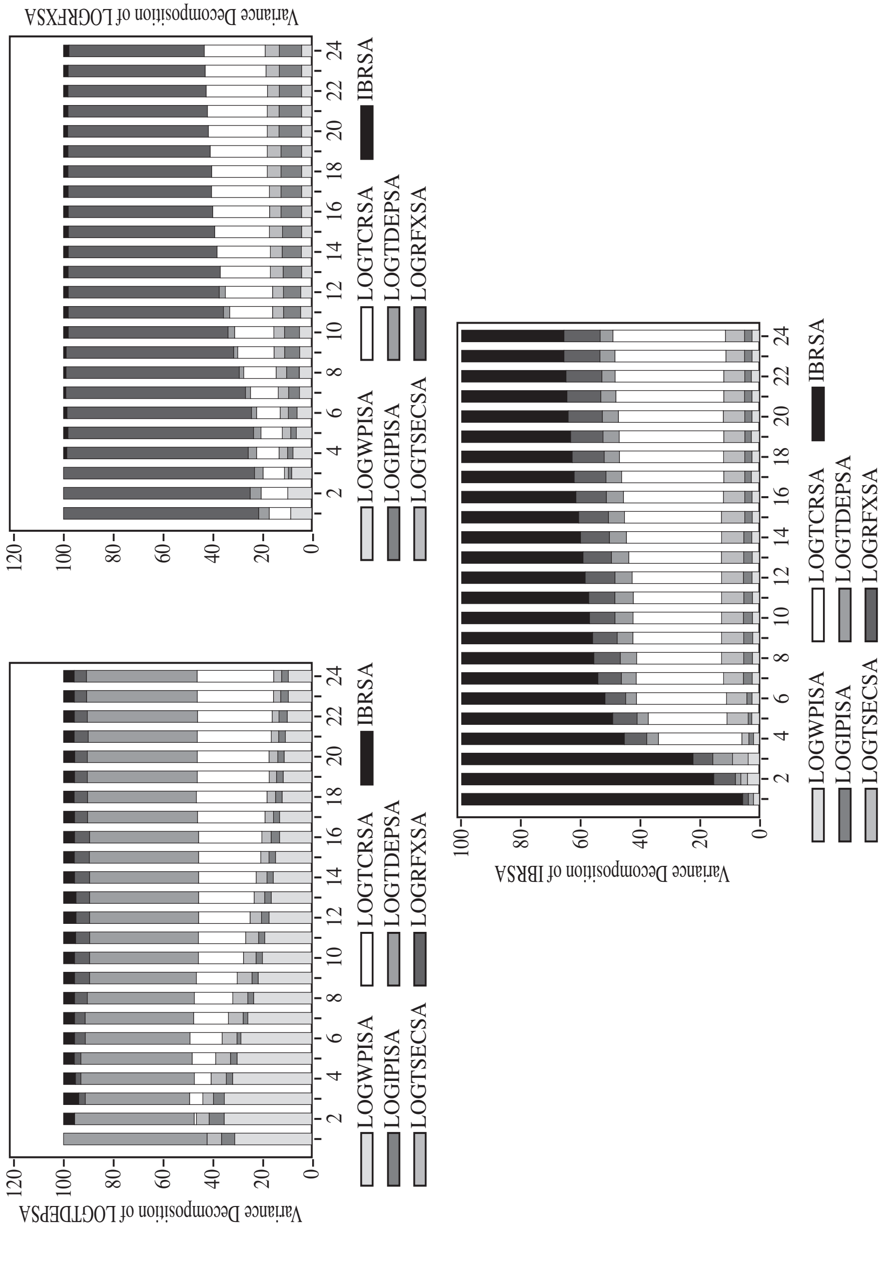
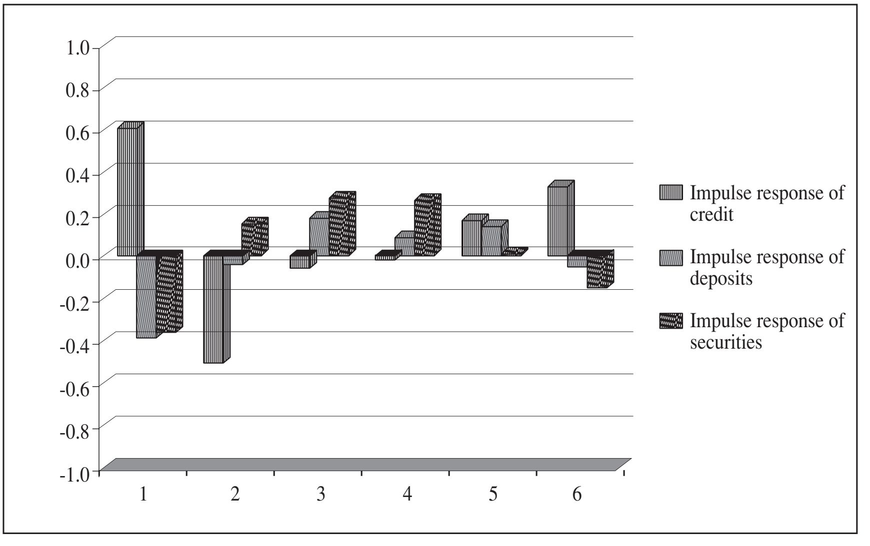
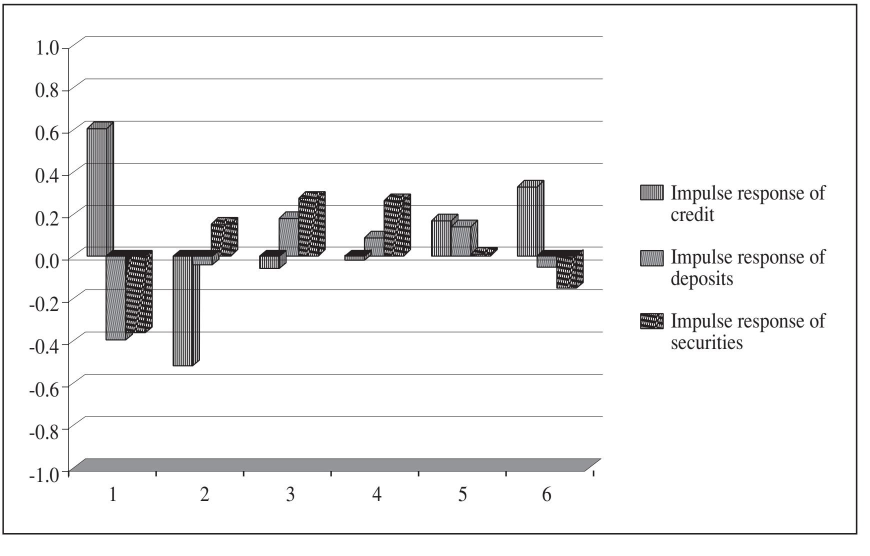


Figure 4: Information Content of Impulse Responses of Financial Variables With Respect to Future Inflation and Future Production

A. Cross Correlation Coefficients vis-a-vis Impulse Responses of Future Industrial Production



B. Cross Correlation Coefficients vis-a-vis Impulse Responses of Future Inflation



Another caveat is in order. It is quite possible that components of total bank credit (e.g., distinguished according to borrower types) would show divergent reactions to an interest rate shock. The idea that the monetary policy has asymmetric effects on certain type of borrowers is in the core of credit view. That is, the credit view encompasses distributional consequences of policy actions, because the costs of finance respond differently for different types of borrowers. It must be pointed out that all evidence presented here refers only to aggregate data. This leaves open the possibility that the credit channel is particularly strong for certain groups of firms, groups of banks or sectors of the economy. These effects may easily be masked by use of aggregate data.

V. Conclusion

In this study, the results of a VAR-analysis of monetary transmission in Turkey have been presented for the period 1986-1998. Especially the relevance of a bank lending channel of monetary policy transmission was investigated at the aggregate level. Since the model could not incorporate disaggregate data reflecting possible differences between the household sector and the corporate sector, the fact that the results are modest should be kept in mind.

It was seen that following a monetary contraction, aggregate bank credit and securities holdings of the banks decline immediately much more than the money (deposits) does. Their somewhat different responses also imply some sort of imperfect substitution between the two assets. Although the identification problem still persists, the timing of impulse responses of the credits and output, and the results of variance decompositions seem to favor bank lending view. Although credit volume as a forecasting variable for the future economic activity performs better than the securities and the deposits, this is inadequate given the high degree of currency substitution and the increasing number of financial instruments. Moreover our results are also consistent with the traditional interest rate channel and the exchange rate channel.

It should be remembered that the credit view is not an alternative to the traditional interest rate mechanism. Its role in the transmission is complementary, as it is the case for exchange rate channel. However their relative importance will change as the economy evolves. One of the implications of credit view is that bank credit supply will offer additional, at times more accurate information about the economy than money. On the other hand given the diverse sectoral responses and differences in access to alternative sources of finance, aggregate bank credit volumes may also be misleading indicators.

This analysis indicates a possible need for broader statistical reporting of some features, including bank loan rates, associated non-price terms, the degree of quantity rationing and volumes of credit, all by sector notably sectors such as households and small businesses which have no alternative source of funds to banks.

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DESEASONALIZING MACROECONOMIC DATA: A CAVEAT TO APPLIED RESEARCHERS IN TURKEY

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Abstract
This paper analyzes the effects of regular seasonal fluctuations of macroeconomic variables in Turkey due to the religious events (religious holidays and Ramadan)[1] in monthly frequency. Conventional deterministic deseasonalization techniques are applied to the detrended and linearized major macroeconomic series. Investigation of the seasonally filtered series reveals residual seasonal regularities vis-à-vis the religious holidays and Ramadan for some of the series. Consequences of ignoring this type of seasonality are also scrutinized.

I. Introduction

Many economic time series exhibit regular seasonal fluctuations. Weather is one reason for such fluctuations in sectors such as tourism, agriculture, and construction. The existence of a fiscal year also causes regular seasonal variation in such variables as government expenditures, car and real estate purchases[2] and the interest rates. The New Year, Mothers' Day,

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1 According to the Hegirian Calendar.

2 Transaction fees are determined once a year, and are put into effect at the beginning of each year. In a high inflation environment, this discrete increase in fees is substantial and causes an increase in purchases of cars and real estate towards the end of the year and a decrease of sales, once the increase in fees actually takes place.

Fathers' Day and the school calendar also have pronounced influence on the retail trade. All these factors, which cause seasonality in economic time series, have fixed dates according to the Gregorian calendar and their effects on these series can be identified without any difficulty using conventional deseasonalization techniques.

While doing time series econometrics analyses, identifying and eliminating regular seasonal fluctuations from each variable entering the estimation, would increase the precision of the coefficient estimates. This is due to two reasons, first, seasonal regularities impose additional variation on variables used in the estimations and second, such fluctuations, in general, are not identical across the dependent and the independent variables.

The method that should be used to eliminate these seasonal fluctuations and the consequences of using different deseasonalization methods on the time series properties of a variable have been studied extensively in the literature.[3]

The purpose of this study is to address the deterministic seasonality issue due to religious holidays in Turkey.[3] Conventional methods of deseasonalization that are suitable for the Gregorian calendar will not detect seasonality of the religious holidays and Ramadan or moving holidays in general that have fixed dates[5] since these events move approximately 11 days earlier every Gregorian year.

Evey year, three significant religious events take place: The holy month of Ramadan,[6] the Feast of Ramadan, and the Feast of Sacrifice. The feast of Ramadan[7] lasts for 3.5 days following the end of the month of Ramadan and the feast of Sacrifice[8] lasts for 4.5 days.[9] Sometimes when

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3 See Lovell (1963), Jorgenson (1964), Grether and Nerlove (1970), Gersovitz and MacKinnon (1978), Barsky and Miron (1989), Jaeger and Kunst (1990), and a more recent survey by Hylleberg (1992a).

4 Turkey has been following the Gregorian calendar according to Law No. 698 passed in December 26, 1925.

5 According to the Lunar or the Hegirian calendar. The lunar calendar is called the Hegirian. It is based on cycles of the moon around the earth while the Gregorian calendar is based on the cycles of the earth around the sun.

6 Ramadan is a month of ritual fasting. Ramadan occurs during the ninth month of the Hegirian calendar.

7 Also referred to as Eid ul Fitr.

8 Also referred to as Eid ul Adha.

9 The official durations of the holidays are decreed by Law No. 2429, article 1B. With respect to the duration, the feasts of Ramadan and Sacrifice are referred to as the lesser and the greater feasts, respectively.

the feast of Ramadan or Sacrifice happens in the middle of the week, the Turkish government decrees the remaining days of the week as a holiday.

There are basically two questions that will be addressed in the paper. What may be evident in the weekly data may average out in the monthly frequency.[10] Hence, first, whether these aforementioned holidays cause regular, seasonal, deterministic fluctuations in the monthly main macroeconomic indicators in Turkey is probed for. The data set is first linearized, seasonally adjusted and detrended using conventional methods. Dummy variables for each of the "moving holidays" are used to detect the existence of any remaining deterministic seasonal patterns. The second question that the paper addresses regards the consequences, if any, of ignoring this type of deterministic seasonality. Analysis on the "conventionally" seasonally adjusted series and "further" adjusted series is done to check for changes in persistence of the series as well as cyclical properties by analyzing autocorrelations and cross-correlations with output.

Our analyses reveal the existence of residual seasonality due to religious holidays in the "conventionally" seasonally adjusted series for a number of key variables including the industrial production index, imports and reserve money. We found that the persistence of the variables containing "residual" seasonality tend to increase with the removal of the deterministic seasonality tied to the lunar calendar. Moreover, the volatility for almost all variables decreases after the removal of residual seasonality. When the cross-correlations with output (industrial production index) were analyzed, we found that the relationship between the variables that contained residual seasonality is weakened implying the existence of a spurious relationship caused by common residual seasonality. For all the other variables, on the other hand, removal of residual seasonality increased the cross-correlations, strengthening our conclusions for the importance of paying special attention to such residual deterministic seasonality.

Section two gives a brief description of the methodology. Section three presents the data and the estimation results. Section four concludes.

[10] Preliminary analyses, using weekly data, reveal that the effects of the religious holidays and Ramadan are significant. However, with quarterly data this effect is not evident.

II. Methodology

Traditional univariate methods of analyzing economic time series are mainly concerned with decomposing the variation in a particular series into trend, seasonal, cyclical and irregular components. The decomposition method for a series is not unique and certain systematic assumptions about the nature of and the interaction among the trend, seasonal, cyclical and irregular components are needed to identify the series. For example, the seasonal component may be deterministic/stochastic or multiplicative/additive in nature. Since there are no guidelines from the microeconomic theory about the functional forms of the aggregate series, we follow the standard practice of the real business cycle literature and assume separable trend and seasonality once the series is linearized. We start out by taking the natural logarithm of the series and then detrend and deseasonalize the series in succession for further analysis. Our ultimate aim is to analyze the cyclical and irregular components of the series for the existence of any residual deterministic seasonality. Our claim is that standard methods of deseasonalization are unable to remove certain deterministic seasonality that evolves through time[11].

Let Y_t be a series of interest. We wish to remove the trend and then the deterministic seasonal component of the series.[12] We employ the spline function proposed by Hodrick and Prescott (1997) that extracts the long-run component of the In Y_t series, τ_t, leaving (In Y_t - τ_t) stationary up to the fourth order. The trend component is chosen to minimize the following quadratic expression:

$$\sum_{t=1}^{T}(\text{In } Y_t - \tau_t)^2 + 14{,}400 \sum_{t=2}^{T} [(\tau_{t+1} - \tau_t) - (\tau_t - \tau_{t-1})]^2$$

and the detrended variable is equal to the difference between In Y_t and τ_t. The filter proposed by Hodrick and Prescott allows the trend component to change slowly across time.[13]

[11] Across the Hegirian Calendar.

[12] Stochastic seasonality is not the focal point of this paper. For stochastic seasonality, see for example, Barsky and Miron (1989) and Hylleberg et al. (1990)

[13] The Hodrick Prescott filter has been subject to criticisms, see for example, King and Rebelo (1993), and Cogley and Nason (1995). However, previous research on the Turkish data by Alper (1998) reveals insignificant differences in results when an alternative filter is considered.

Next, we carry out the seasonal adjustment of the trend-free series by estimating its seasonal deterministic component and then removing this component from the trend-free series. To remove the deterministic seasonal component, we use the regression method due to Lovell (1963) and Jorgenson (1964)[14] and estimate the following model:

$$(\text{In } Y_t - \tau_t) = \sum_{i=1}^{12} \alpha_i D_{it} + \sum_{j=1}^{s} \beta_j P_{jt} + u_t$$

where u_t is a stochastic component that may or may not be white noise, D_{it}, i=1, ..., 12, are monthly dummies and P_{jt}, j=1, ..., s are polynomial terms in time up to order $s \geq 1$. The latter variables are included to account for the non-seasonal deterministic component. We get the "conventionally" seasonally adjusted variable, c_t, as

$$c_t = (\text{In } Y_t - \tau_t) - \sum_{i=1}^{12} \hat{\alpha}_i - D_{it}$$

We suspect that c_t still contains some deterministic seasonality, that is, regular seasonal peaks and troughs, which still exists due to the moving holidays tied to the lunar calendar. For the detection of residual seasonality due to moving holidays, we estimate the equation

$$c_t = \sum_{i=1}^{4} \delta_i d_{it} + \sum_{k=1}^{r} \phi_k c_{t-k} + \varepsilon_t$$

where ε_t is a stochastic component that is serially uncorrelated, d_{it} is a monthly seasonal dummy variable that takes the value 1 if a religious holiday or Ramadan tied to the lunar calendar takes place that particular month, zero otherwise. Initially, ignoring the religious dummy variables, we identify r, the order of the autoregressive process at the right hand side of the equation, by choosing the minimum value making ε_t is serially uncorrelated. We then estimate the autoregressive process including the religious intercept dummy variables and check for the significance of the

[14] The X-11 method of the U.S. Bureau of Census, which is a variant of the moving average method, is also used as an alternative method to deseasonalize the series when possible stochastic seasonality is present. See Hylleberg (1992b) for the details of this method. The results turned out to be quite similar. We chose the regression method over X-11 due to the loss of reliability at the end series as well as the 'excess persistence' findings by Jaeger and Kunst (1990) of the X-11 adjusted data compared to data adjusted by regression on dummies.

dummy variables. Significant coefficient(s) of the dummy variables is an indication of "leftover" or "residual" deterministic seasonality, since with the removal of trend and seasonality and a reasonably well-fit autoregressive process, what remains should be a pure random component, not explained by any variable.

Next, by using the estimates from the regression above, we further deseasonalize the series by estimating the following equation:

$$c_t = \sum_{i=1}^{4} \gamma_i d_{it} + \sum_{j=1}^{s} \beta_j P_{jt} + u_t$$

and then subtracting the effects of the dummy variables from c_t.

$$f_t = c_t - \sum_{i=1}^{4} \hat{\gamma}_i d_{it}$$

We next analyze the consequences of ignoring this "residual" seasonality. As mentioned previously, improperly idendifying and eliminating regular seasonal fluctuations from variables used in time series analyses reduce the precision of the coefficient estimates since seasonal regularities impose additional variation on variables used in the estimations. For a number of macroeconomic monthly time series, we calculate the autocorrelation functions and check whether or not persistence increases since there exists less noise in the data once the deterministic " Seasonality" is eliminated. We also check to see whether the volatility of each series reduces once the leftover seasonality is removed. Finally, we calculate monthly cross-correlations and look for any emerging patterns after the residual deterministic seasonality is eliminated.

III. Data and Empirical Results

The monthly data set used in the paper, covering the period January 1985-August 2000, consists of 24 variables and is obtained from the internet site of the Central Bank of the Republic of Turkey (www.tcmb.gov.tr). Table 1 gives the definitions and the ranges of each series used in the analyses.

As explained in the methodology, we first take the natural logarithm of these variables, then obtain the trend-free series using the Hodrick-Prescott filter and finally deseasonalize the series using the method due to Lovell (1963) and Jorgenson (1964). The resulting series is considered to be the irregular component of the variable with no trend and no regular seasonal fluctuations. Next, we identify the order of the autoregressive process for each of the 24 trend-free, deseasonalized and linearized series. The order of these processes are chosen based on two criteria: first the residuals from the estimation must be serially uncorrelated, and second, the principle of parsimony. After the autoregressive order of each series is identified, dummy variables representing the moving holidays are appended to the estimation; and based on Schwarz criterion (1978) and Wald tests of coefficient restrictions, the significance of these dummy variables in the regression are tested.

Four dummy variables are created to represent the three religious feast holidays and Ramadan.[15] The first and the second dummy variables represent the feasts of Ramadan and Sacrifice and take on the value 1 for a month if that month contains at least half of the respective feasts (2 and 2.5 days, respectively) and zero otherwise. The third dummy variable is for the extended holiday, and it takes on the value 1 if the government has extended the holiday to five business days for the feast and zero otherwise. The fourth dummy variable is for the Holy month of Ramadan and takes on the value 1 if a month contains at least 5 business days of it, zero otherwise. Thus, while the first three dummy variables cannot take on the value 1 for two consecutive "Gregorian" months, this is not necessarily true for the Ramadan dummy. The values of the four dummy variables are presented in Table 2.

Before turning to formal estimations, Figure 1 provides informal evidence of the existence of residual seasonality. It plots the detrended, deseasonalized industrial production index before and after the removal of deterministic residual seasonality. Some of the spikes (troughs and peaks) in the data disappear (e.g. 1987, 1990, 1991, 1997) once the religious events are controlled for.

Table 3 summarizes the results for the formal detection of residual seasonality. First four columns give the order of the autoregressive process, the Q-statistic[16] for testing the existence of serial correlation in the residuals up to 24 lags, the adjusted R-squared and the Schwarz criterion for each regression. All the reported Q-statistics lead to failure of rejection of the null hypothesis, implying that the residuals from the autoregressive models are serially uncorrelated. The fifth column gives information

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[15] The exact dates of these events for the post 1987 period are obtained from the Directorate of Religious Affairs of Turkey.

[16] The Q-statistic, whose null hypothesis is no autocorrelation up to a pre-specified number of lags, is due to Ljung and Box (1979). It asymptotically follows a chi-square distribution with degrees of freedom equal to the number of lags.

about the significance of the religious dummy variables, when they are included in the “identified” autoregressive regression. A minus sign indicates insignificance of the coefficients for all religious dummy variables. After all the dummy variables are added to the regression, the variables with insignificant coefficients are taken out and the test statistics for the remaining variables are reported. The values of the coefficients as well as the corresponding p-values for significance are given in the fifth column. As mentioned earlier, if the dummy variables are significant, it would imply the existence of “unaccounted-for seasonality” in the data. Of the 24 variables examined, 9 variables display significant effects of at least one of the religious holidays and Ramadan. The next column reports the Wald test statistics[17] for testing the joint significance of the coefficients of the dummy variables of the religious holidays and Ramadan and the corresponding p-values. The last two columns report the adjusted R-squared and the Schwarz criterion associated with the regression including the dummy variables.

For variables like the industrial production index and its subgroups, the coefficients of the corresponding dummy variables are significantly negative due to, the loss of business days. We also observe that reserve money increases significantly for the months having the two feasts, implying that the open market operations by the central bank provide liquidity to the market during the holidays. These operations are carried out in response to an increase in the liquidity demand prior to the holidays. Moreover, government expenditures increase significantly during the month of Ramadan. It is worthwhile to note that Ramadan and the extended holiday dummy variables are significant only one occasion but the feast of Ramadan and feast of Sacrifice have significant effects for almost all 9 variables.

After verifying the existence of residual deterministic seasonality[18], we turn to possible consequences of ignoring these effects. For this purpose, we obtain cross-correlations and auto-correlation tables of the detrended and deseasonalized variables with and without the residual seasonality. Table 4 reports the autocorrelations up to six lags for variables with sig-

[17] The Wald statistic asymptotically follows an F distribution with q,(n-k) degrees of freedom where q is the number of restrictions, n is the sample of variables and k is the number of independent variables in the regression. Equivalently, a chi-square distribution could have also been used. All our results based on the F distribution-based Wald test are also obtained by the chi-square distribution-based Wald test.

[18] Due to the existence of the implicit Hegirian Calendar effects.

nificant residual seasonality as found in Table 3, except central bank money and credits since these variables show relatively less significant residual seasonality given by the p-values of the corresponding Wald tests. The upper half of the rows report results pertaining to the linearized, detrended, and deseasonalized data containing "residual" seasonality, the lower half of the rows report autocorrelations after the "residual" seasonality is removed. Conforming to the a priori expectations, almost all autocorrelation coefficients, giving information about the persistence[19] of the data, rose once the residual seasonality is removed. This increase was about 4 percentage points on average. This implies that after the removal of residual seasonality, the lags of these variables had increased predictive power for explaining the current level.

Next, as is standard in the business cycle literature, cross correlations of some of the series[20] with the Industrial Production Index and their volatilities are analyzed. The results are reported in Table 5. Again, conforming to the a priori expectations, the volatility of most of the series[21] reduced once the noise from the "residual seasonality" is removed. Correlation coefficients that are greater than 0.20 in absolute value are boldfaced to imply statistical significance. When the cross correlation coefficients prior to the removal of the deterministic residual seasonality are compared to those obtained after the removal two important trends emerge. First, for the variables, which had significant "residual seasonality", many of the coefficients in the lower part of the table are less than their counterparts in the upper part of the table in absolute value. On the other hand, for the variables, which did not show any significant residual seasonality, only five coefficients[22] out of 55 significant coefficients are less than their counterparts in the upper part of the table in absolute value. These results imply that for the variables having significant residual seasonality, the cross-correlation coefficients with industrial production index are overstated since they capture the co-movement of the "residual"

[19] Informally, persistence may be defined as the long-run level effect of a 1 per cent shock on a macroeconomic time series.

[20] It should be noted that the series also include variables with no deterministic residual seasonality. The variables omitted in this table yield insignificant correlation coefficients with the industrial production index.

[21] The volatility of all series except foreign exchange denominated deposits and the central bank money decreases significantly. Volatility increases for these series are statistically insignificant.

[22] These are the third lag of credits, stock exchange index in TL and in USD, second lag of M1 and first lag of CPI inflation.

seasonality in the series and industrial production index. For all other variables, the cross-correlations increase in absolute value, once the "residual" seasonality is removed. Even though these variables do not have any residual seasonality, we still get an increase due to the removal of seasonality in industrial production index.

To recapitulate, while the persistence of the series with "residual" seasonality increase after the removal of the latter, volatility of all series is reduced. These two results imply that after the removal of residual seasonality, the series become more predictable and estimation results based on these variables will be more reliable. Moreover, the correlations with Industrial Production Index decrease in absolute value for the variables with "residual" seasonality, signifying the existence of a spurious relationship due to "residual" seasonality whereas for all others they increase in absolute value, strengthening the conclusions above.

IV. Conclusion

Proper decomposition of a macroeconomic time series into a trend, seasonal, cyclical and irregular components is essential for an econometrician to make inferences about the unknown population parameters that are of interest to economic theory. The aim of this study is to show that for Turkey, conventional deseasonalization procedures, such as regression on dummy variables or X-11, may fail to remove all deterministic seasonality when certain significant events, such as religious holidays, follow a different calendar. Turkey has been following the Gregorian calendar, based on the cycles of the earth around the sun, since 1926. However, religious feasts and Ramadan are based on the lunar cycles. The impact of these events on economic variables may escape the detection of conventional deseasonalization methods that search for regular monthly peaks and troughs in the series.

We analyze 24 monthly macroeconomic time series for Turkey, by first linearizing the series and then obtaining the trend-free, deterministic seasonal-free variables using conventional methods. Of the 24 variables examined, 9 variables contain significant effects of at least one of the three religious events. These 9 variables include measures of aggregate economic activity such as the industrial and the manufacturing production indexes, monetary measures such as reserve money and government revenues. Price-like variables and financial variables do not show significant signs of residual seasonality.

Upon detecting the existence of residual deterministic seasonality in a number of series, a search has been made for the consequences of ignoring such seasonality. We calculated the autocorrelations, volatilities and cross-correlations with output for the variables before and after the removal of the "residual" seasonality due to moving holidays. We found that the persistence of the variables containing "residual" seasonality tend to increase with the removal of the deterministic seasonality tied to the lunar calendar. Moreover, the volatility for almost all variables decreases after the removal of residual seasonality. These two facts imply that the variables become relatively more predictable and estimation results based on these variables will be more reliable when the residual seasonality is removed. When the cross-correlations with output (industrial production index) were analyzed, we found that the relationship between the variables that contained residual seasonality is weakened implying the existence of a spurious relationship caused by common residual seasonality. For all the other variables, on the other hand, removal of residual seasonality increased the cross-correlations, strengthening our conclusions for the importance of paying special attention to such residual deterministic seasonality.

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Figure 1: Detrended Industrial Production Index Before and After the Removal of Residual Seasonality

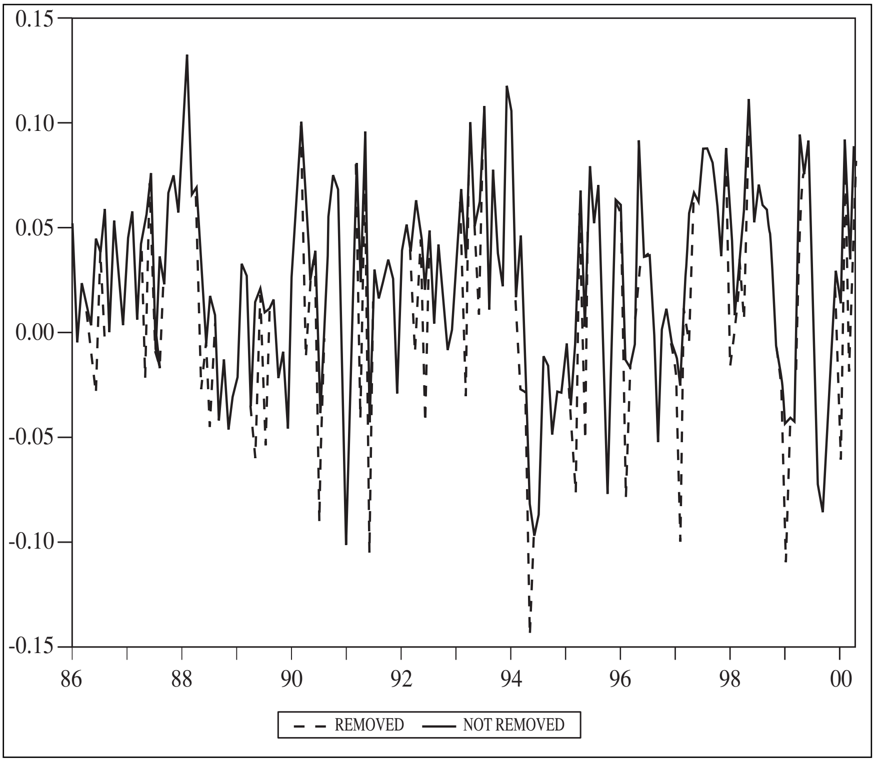

Table 1: Descriptions and Ranges of Variables

Acronym	Description	Starting Date	Ending Date	Number of Observations
IPROD	SIS Industrial Production Index (1992=100)	Jan 1986	April 2000	172
IPIMQ	Mining and Quarrying (subgroup of IPROD)	Jan 1986	April 2000	172
MPROD	Manufacturing Industries (subgroup of IPROD)	Jan 1986	April 2000	172
IPEGW	Electricity Gas and Water (subgroup of IPROD)	Jan 1986	April 2000	172
IMPOR	Imports (in million USD)	Jan 1985	May 2000	185
GOVRE	Government Revenues (in million TRL)	Jan 1985	June 2000	186
GOVEX	Government Expenditures (in million TRL)	Jan 1985	June 2000	186
NDOMB	Net Domestic Borrowing (in million TRL)	Jan 1985	June 2000	186
M1	M1 (in billion TRL)	Jan 1986	April 2000	172
FXDEP	Foreign Exchange Denominated Deposit Accounts (in billion TRL)	Jan 1986	April 2000	172
CBM	Central Bank Money (in million TRL)	Sep 1989	Aug 2000	132
RM	Reserve Money (in million TRL)	Sep 1989	Aug 2000	132
WPI	SIS Whole Sale Price Index (1987=100)	Jan 1985	June 2000	186
CPI	SIS Consumer Price Index (1987=100)	Jan 1987	June 2000	162
WPIINF	WPI Inflation (year-on-year)	Jan 1986	June 2000	174
CPIINF	CPI Inflation (year-on-year)	Jan 1988	June 2000	150
CREDIT	Credits given by Deposit Banks (in billion TRL)	Jan 1986	April 2000	172
ISETL	Istanbul Stock Exchange National-100 Index (Monthly Average, TRL Based)	Jan 1986	Aug 2000	176
ISEVOL	Istanbul Stock Exchange Trade Volume (Monthly Average)	Nov 1986	Aug 2000	166
ISEUSD	Istanbul Stock Exchange National-100 Index (Monthly Average, USD Based)	Jan 1986	Aug 2000	176
ISEFIN	Istanbul Stock Exchange Financial-Index (Monthly Average, USD Based)	Jan 1991	Aug 2000	116
ISEIND	Istanbul Stock Exchange Industrial Index (Monthly Average, USD Based)	Jan 1991	Aug 2000	116
ONINTR	Weighted Average of Overnight Simple Interest Rate in Interbank Market	Jan 1990	Aug 2000	128
USDTL	Exchange Rate of USD (Central Bank Buying Rate)	Jan 1985	Aug 2000	188

Table 2: The Values of the Dummy Variables

Month	Feast of Ramadan	Feast of Sacrifice	Nine Days	Ramadan	Month	Feast of Ramadan	Feast of Sacrifice	Nine Days	Ramadan
Jan-85	0	0	0	0	Dec-88	0	0	0	0
Feb-85	0	0	0	0	Jan-89	0	0	0	0
Mar-85	0	0	0	0	Feb-89	0	0	0	0
Apr-85	0	0	0	0	Mar-89	0	0	0	0
May-85	0	0	0	1	Apr-89	0	0	0	1
Jun-85	1	0	0	1	May-89	1	0	0	1
Jul-85	0	0	0	0	Jun-89	0	0	0	0
Aug-85	0	1	1	0	Jul-89	0	1	0	0
Sep-85	0	0	0	0	Aug-89	0	0	0	0
Oct-85	0	0	0	0	Sep-89	0	0	0	0
Nov-85	0	0	0	0	Oct-89	0	0	0	0
Dec-85	0	0	0	0	Nov-89	0	0	0	0
Jan-86	0	0	0	0	Dec-89	0	0	0	0
Feb-86	0	0	0	0	Jan-90	0	0	0	0
Mar-86	0	0	0	0	Feb-90	0	0	0	0
Apr-86	0	0	0	0	Mar-90	0	0	0	0
May-86	0	0	0	1	Apr-90	1	0	0	1
Jun-86	1	0	0	1	May-90	0	0	0	0
Jul-86	0	0	0	0	Jun-90	0	0	0	0
Aug-86	0	1	0	0	Jul-90	0	1	1	0
Sep-86	0	0	0	0	Aug-90	0	0	0	0
Oct-86	0	0	0	0	Sep-90	0	0	0	0
Nov-86	0	0	0	0	Oct-90	0	0	0	0
Dec-86	0	0	0	0	Nov-90	0	0	0	0
Jan-87	0	0	0	0	Dec-90	0	0	0	0
Feb-87	0	0	0	0	Jan-91	0	0	0	0
Mar-87	0	0	0	0	Feb-91	0	0	0	0
Apr-87	0	0	0	0	Mar-91	0	0	0	1
May-87	1	0	0	1	Apr-91	1	0	1	1
Jun-87	0	0	0	0	May-91	0	0	0	0
Jul-87	0	0	0	0	Jun-91	0	1	0	0
Aug-87	0	1	1	0	Jul-91	0	0	0	0
Sep-87	0	0	0	0	Aug-91	0	0	0	0
Oct-87	0	0	0	0	Sep-91	0	0	0	0
Nov-87	0	0	0	0	Oct-91	0	0	0	0
Dec-87	0	0	0	0	Nov-91	0	0	0	0
Jan-88	0	0	0	0	Dec-91	0	0	0	0
Feb-88	0	0	0	0	Jan-92	0	0	0	0
Mar-88	0	0	0	0	Feb-92	0	0	0	0
Apr-88	0	0	0	1	Mar-92	0	0	0	1
May-88	1	0	1	1	Apr-92	1	0	0	1
Jun-88	0	0	0	0	May-92	0	0	0	0
Jul-88	0	1	0	0	Jun-92	0	1	0	0
Aug-88	0	0	0	0	Jul-92	0	0	0	0
Sep-88	0	0	0	0	Aug-92	0	0	0	0
Oct-88	0	0	0	0	Sep-92	0	0	0	0
Nov-88	0	0	0	0	Oct-92	0	0	0	0

Table 2: The Values of the Dummy Variables (continued)

Month	Feast of Ramadan	Feast of Sacrifice	Nine Days	Ramadan	Month	Feast of Ramadan	Feast of Sacrifice	Nine Days	Ramadan
Nov-92	0	0	0	0	Oct-96	0	0	0	0
Dec-92	0	0	0	0	Nov-96	0	0	0	0
Jan-93	0	0	0	0	Dec-96	0	0	0	0
Feb-93	0	0	0	1	Jan-97	0	0	0	1
Mar-93	1	0	1	1	Feb-97	1	0	0	1
Apr-93	0	0	0	0	Mar-97	0	0	0	0
May-93	0	0	0	0	Apr-97	0	1	0	0
Jun-93	0	1	1	0	May-97	0	0	0	0
Jul-93	0	0	0	0	Jun-97	0	0	0	0
Aug-93	0	0	0	0	Jul-97	0	0	0	0
Sep-93	0	0	0	0	Aug-97	0	0	0	0
Oct-93	0	0	0	0	Sep-97	0	0	0	0
Nov-93	0	0	0	0	Oct-97	0	0	0	0
Dec-93	0	0	0	0	Nov-97	0	0	0	0
Jan-94	0	0	0	0	Dec-97	0	0	0	0
Feb-94	0	0	0	1	Jan-98	1	0	0	1
Mar-94	1	0	0	1	Feb-98	0	0	0	0
Apr-94	0	0	0	0	Mar-98	0	0	0	0
May-94	0	1	0	0	Apr-98	0	1	1	0
Jun-94	0	0	0	0	May-98	0	0	0	0
Jul-94	0	0	0	0	Jun-98	0	0	0	0
Aug-94	0	0	0	0	Jul-98	0	0	0	0
Sep-94	0	0	0	0	Aug-98	0	0	0	0
Oct-94	0	0	0	0	Sep-98	0	0	0	0
Nov-94	0	0	0	0	Oct-98	0	0	0	0
Dec-94	0	0	0	0	Nov-98	0	0	0	0
Jan-95	0	0	0	0	Dec-98	0	0	0	1
Feb-95	0	0	0	1	Jan-99	1	0	1	1
Mar-95	1	0	0	1	Feb-99	0	0	0	0
Apr-95	0	0	0	0	Mar-99	0	0	0	0
May-95	0	1	1	0	Apr-99	0	1	1	0
Jun-95	0	0	0	0	May-99	0	0	0	0
Jul-95	0	0	0	0	Jun-99	0	0	0	0
Aug-95	0	0	0	0	Jul-99	0	0	0	0
Sep-95	0	0	0	0	Aug-99	0	0	0	0
Oct-95	0	0	0	0	Sep-99	0	0	0	0
Nov-95	0	0	0	0	Oct-99	0	0	0	0
Dec-95	0	0	0	0	Nov-99	0	0	0	0
Jan-96	0	0	0	1	Dec-99	0	0	0	1
Feb-96	1	0	1	1	Jan-00	1	0	0	1
Mar-96	0	0	0	0	Feb-00	0	0	0	0
Apr-96	0	0	0	0	Mar-00	0	1	1	0
May-96	0	1	0	0	Apr-00	0	0	0	0
Jun-96	0	0	0	0	May-00	0	0	0	0
Jul-96	0	0	0	0	Jun-00	0	0	0	0
Aug-96	0	0	0	0	Jul-00	0	0	0	0
Sep-96	0	0	0	0	Aug-00	0	0	0	0

Table 3: Regression Results

Variable	Fitted Model	Q Statistic for 24 lags	Adjusted R-squared and SC for the Original Model		Religious Events Event / Coefficient / (P-Value)			Wald Test Statistic	Adjusted R-squared and SC for the Model with Dummies	
IPROD	AR(13)	17.47 (0.83)	0.26	-3.21	Feast of Ramadan Feast of Sacrifice	-0.04 -0.05	(0.00) (0.00)	11.48 (0.00)	0.35	-3.30
IPIMQ	AR(13)	17.87 (0.81)	0.18	-2.43	Feast of Sacrifice	-0.04	(0.03)	4.68 (0.03)	0.20	-2.43
MPROD	AR(13)	16.71 (0.86)	0.26	-2.95	Feast of Ramadan Feast of Sacrifice	-0.04 -0.05	(0.00) (0.00)	10.62 (0.00)	0.35	-3.02
IPEGW	AR(13)	17.10 (0.84)	0.09	-4.16	Feast of Ramadan Feast of Sacrifice	-0.03 -0.03	(0.00) (0.00)	16.81 (0.00)	0.25	-4.30
IMPOR	AR(11)	20.57 (0.66)	0.51	-1.69	Feast of Ramadan Feast of Sacrifice	-0.10 -0.07	(0.00) (0.00)	12.60 (0.00)	0.57	-1.78
GOVRE	AR(14)	21.37 (0.62)	0.67	-3.35	Feast of Ramadan	-0.02	(0.03)	4.84 (0.03)	0.67	-3.54
GOVEX	AR(13)	24.24 (0.45)	0.30	-2.43		-		-	-	-
NDOMB	AR(4)	26.88 (0.31)	0.43	0.13		-		-	-	-
RM	AR(13)	20.68 (0.66)	0.53	-4.06	Feast of Ramadan Feast of Sacrifice	0.02 0.02	(0.04) (0.02)	4.22 (0.02)	0.55	-4.06
M1	AR(10)	21.96 (0.58)	0.45	-3.06		-		-	-	-
FXDEP	AR(8)	16.30 (0.88)	0.74	-3.33		-		-	-	-
CBM	AR(12)	11.48 (0.99)	0.74	-1.76	Nine days	0.06	(0.04)	4.12 (0.04)	0.75	-1.76
WPI	AR(2)	22.63 (0.54)	0.86	-4.87		-		-	-	-
CPI	AR(6)	17.59 (0.82)	0.81	-7.98		-		-	-	-
WPIINF	AR(13)	11.25 (0.99)	0.88	-3.15		-		-	-	-
CPIINF	AR(10)	24.59 (0.43)	0.83	-4.00		-		-	-	-
CREDIT	AR(4)	31.09 (0.15)	0.89	-4.66	Feast of Ramadan Ramadan	0.02 -0.05	(0.03) (0.03)	2.81 (0.06)	0.89	-4.64

Table 3: Regression Results (continued)

Variable	Fitted Model	Q Statistic for 24 lags	Adjusted R-squared and SC for the Original Model		Religious Events Event / Coefficient / (P-Value)			Wald Test Statistic	Adjusted R-squared and SC for the Model with Dummies	
ISETL	AR(11)	12.96 (0.97)	0.88	-1.01		-		-	-	-
ISEVOL	AR(12)	19.10 (0.75)	0.70	1.03		-		-	-	-
ISEUSD	AR(11)	9.68 (0.99)	0.89	-0.93		-		-	-	-
ISEFIN	AR(11)	9.22 (0.99)	0.80	-0.59		-		-	-	-
ISEIND	AR(11)	14.77 (0.93)	0.80	-1.06		-		-	-	-
ONINTR	AR(1)	17.52 (0.83)	0.46	-0.42		-		-	-	-
USDTL	AR(13)	6.86 (0.99)	0.86	-3.69		-		-	-	-

Notes: Numbers in parentheses below a test statistic is the p-value for that test statistic. For coefficients, the numbers in parentheses are the p-values for the simple t-statistics. In the Religious Events columns, a dash represents no significant effect of religious dummy variables. All t-statistics and the Wald test statistics are significant with 5% significance (except for CREDIT, the Wald test statistic has a p-value of 0.06%. All Q-statistics are insignificant, the smallest p-value being 0.15.

* Religious holidays and Ramadan.

** When the religious holidays and Ramadan happens in the middle of the week, the Turkish Government decrees the remaining days of the week as a holiday i.e. nine days.

Table 4: Autocorrelations of Chosen Macroeconomic Series Before and After the Removal of the Seasonality Effects

Before Removal	**t-6**	**t-3**	**t-2**	**t-1**	**t**	**t+1**	**t+2**	**t+3**	**t+6**
IPROD	0.01	0.04	**0.26**	**0.28**	**1.00**	**0.28**	**0.26**	0.04	0.01
IPIMQ	**-0.24**	0.05	**0.23**	**0.48**	**1.00**	**0.48**	**0.23**	0.05	**-0.24**
MPROD	0.02	0.03	**0.26**	**0.31**	**1.00**	**0.31**	**0.26**	0.03	0.02
IPEGW	-0.10	-0.09	0.11	**0.24**	**1.00**	**0.24**	0.11	-0.09	-0.10
IMPOR	0.06	**0.42**	**0.61**	**0.56**	**1.00**	**0.56**	**0.61**	**0.42**	0.06
GOVRE	**0.38**	**0.62**	**0.72**	**0.82**	**1.00**	**0.82**	**0.72**	**0.62**	**0.38**
RM	-0.13	**0.26**	**0.53**	**0.68**	**1.00**	**0.68**	**0.53**	**0.26**	-0.13
Seasonality Effects Removed	**t-6**	**t-3**	**t-2**	**t-1**	**t**	**t+1**	**t+2**	**t+3**	**t+6**
IPROD	0.05	0.08	**0.27**	**0.49**	**1.00**	**0.49**	**0.27**	0.08	0.05
IPIMQ	**-0.17**(*)	0.08	**0.20**(*)	**0.50**	**1.00**	**0.50**	**0.20**(*)	0.03	**-0.17**(*)
MPROD	0.06	0.06	**0.27**	**0.49**	**1.00**	**0.49**	**0.27**	0.06	0.06
IPEGW	**-0.27**	0.01	**0.16**	**0.52**	**1.00**	**0.52**	**0.16**	0.01	**-0.27**
IMPOR	0.07	**0.48**	**0.63**	**0.66**	**1.00**	**0.66**	**0.63**	**0.48**	0.07
GOVRE	**0.38**	**0.64**	**0.72**	**0.82**(*)	**1.00**	**0.82**(*)	**0.72**	**0.64**	**0.38**
RM	-0.17	**0.30**	**0.57**	**0.77**	**1.00**	**0.77**	**0.57**	**0.30**	-0.17

Notes: The variables in this table are those which have significant results as in Table 2 except Central Bank money and credits since these variables show relatively less significant residual seasonality given by the p-values of the corresponding Wald tests. Numbers in boldface reflect significant autocorrelation coefficients at 95%. The significance is determined by threshold values (0.1524 for the first four variables, 0.147 for imports and government revenues and 0.174 for reserve money) which are calculated using the aymptotic standard error. ($1/\sqrt{T}$)

(*) sign reflects a decrease in the absolute value compared to the value before the religious seasonality is removed.

Table 5: Cross Correlations of Some Series with the Industrial Production Index, Before and After the Removal of the Seasonality Effects

Before Removal	Volatility (Std. Error in Percentage)	t-6	t-3	t-2	t-1	t	t+1	t+2	t+3	t+6
IPROD	5.28%	0.01	**0.04**	**0.26**	**0.28**	**1.00**	**0.28**	**0.26**	0.04	0.01
IPIMQ	8.12%	-0.14	-0.02	0.06	0.05	0.31	0.02	0.03	-0.03	-0.07
MPROD	6.11%	0.02	0.03	**0.26**	**0.29**	**0.99**	**0.29**	**0.27**	0.04	0.01
IPEGW	3.21%	0.04	0.03	0.16	-0.02	**0.49**	0.02	0.06	0.07	0.05
IMPOR	13.06%	0.02	**0.22**	**0.39**	**0.33**	**0.61**	**0.28**	**0.31**	0.14	-0.09
RM	3.96%	0.02	**0.22**	0.17	0.19	**-0.26**	-0.12	**-0.25**	-0.18	-0.11
WPI	5.45%	-0.01	-0.15	-0.18	**-0.23**	**-0.25**	**-0.25**	**-0.24**	-0.18	-0.14
CPI	4.07%	-0.03	-0.18	**-0.21**	**-0.23**	**-0.20**	-0.16	-0.13	-0.03	-0.05
CREDIT	6.72%	0.07	**0.25**	**0.28**	**0.35**	**0.38**	**0.42**	**0.43**	**0.43**	**0.29**
ISETL	35.95%	0.03	**0.24**	**0.28**	**0.30**	**0.26**	0.19	0.11	0.06	0.01
ISEVOL	65.31%	0.01	**0.27**	**0.31**	**0.32**	**0.29**	0.17	0.10	0.06	-0.01
M1	6.43%	0.11	**0.31**	**0.39**	**0.22**	-0.01	0.09	-0.09	-0.03	-0.12
WPIINF	12.19%	-0.05	-0.20	-0.19	-0.17	-0.15	-0.14	-0.11	-0.09	-0.05
CPIINF	7.01%	0.03	-0.09	-0.16	**-0.23**	**-0.21**	-0.16	-0.12	-0.06	-0.01
ISEUSD	39.26%	0.05	**0.27**	**0.33**	**0.37**	**0.32**	**0.25**	0.17	0.11	0.02
ONINTR	26.87%	-0.10	**-0.36**	**-0.44**	**-0.34**	-0.19	0.03	0.13	0.17	0.14
ISEFIN	33.25%	-0.15	0.20	**0.33**	**0.38**	**0.32**	**0.21**	0.14	0.08	-0.02
ISEIND	26.07%	-0.07	0.23	**0.37**	**0.43**	**0.36**	**0.24**	0.15	0.07	-0.07
USDTL	8.54%	-0.10	**-0.22**	**-0.31**	**-0.39**	**-0.39**	**-0.34**	**-0.31**	**-0.24**	-0.06

Table 5: Cross Correlations of Some Series with the Industrial Production Index, Before and After the Removal of the Seasonality Effects (continued)

Seasonality Effects Removed	**Volatility (Std. Error in Percentage)**	**t-6**	**t-3**	**t-2**	**t-1**	**t**	**t+1**	**t+2**	**t+3**	**t+6**
IPROD	4.65%	0.05	0.08	**0.27**	**0.50**	**1.00**	**0.5**	**0.27**	0.08	0.05
IPIMQ	7.67%	-0.07(*)	-0.02	-0.00(*)	0.05(*)	**0.20**(*)	0.09	-0.02(*)	-0.05	0.00(*)
MPROD	5.44%	0.06	0.07	**0.27**	**0.49**	**0.99**(*)	**0.49**	**0.27**	0.08	0.05
IPEGW	2.94%	0.02(*)	0.11	0.16	**0.24**	**0.43**(*)	0.20	0.10	0.11	0.02(*)
IMPOR	12.48%	0.06	**0.28**	**0.40**	**0.46**	**0.56**(*)	**0.43**	**0.36**	**0.20**	-0.08(*)
RM	3.75%	0.06	0.19(*)	0.17	0.05(*)	-0.19(*)	**-0.24**	**-0.25**	**-0.21**	-0.10(*)
WPI	5.41%	-0.01	-0.15(*)	**-0.22**	**-0.24**	**-0.29**	**-0.27**	**-0.27**	**-0.23**	-0.14
CPI	4.05%	-0.03(*)	-0.16(*)	**-0.25**	**-0.26**	**-0.25**	-0.19	-0.16	-0.07	-0.03(*)
CREDIT	6.68%	0.07	**0.25**(*)	**0.33**	**0.36**	**0.41**	**0.48**	**0.47**	**0.47**	**0.33**
ISETL	35.52%	0.05	**0.23**(*)	**0.31**	**0.33**	**0.31**	**0.22**	0.14	0.08	-0.01
ISEVOL	64.18%	0.02	**0.28**	**0.36**	**0.35**	**0.34**	0.20	0.10	0.08	-0.05
M1	6.38%	0.14	**0.33**	**0.33**(*)	**0.23**	0.01(*)	0.04(*)	-0.01(*)	-0.04	-0.11(*)
WPIINF	11.95%	-0.06	-0.18(*)	**-0.21**	-0.16(*)	-0.17	-0.13(*)	-0.13	-0.12	-0.04(*)
CPIINF	6.92%	-0.01(*)	-0.07(*)	-0.17	**-0.21**(*)	**-0.22**	-0.15(*)	-0.12	-0.08	-0.02
ISEUSD	38.66%	0.07	**0.27**(*)	**0.37**	**0.39**	**0.38**	0.28	0.20	0.14	0.00(*)
ONINTR	26.67%	-0.15	**-0.40**	**-0.46**	**-0.37**	-0.18(*)	0.00(*)	0.15	0.17(*)	0.14(*)
ISEFIN	33.24%	-0.15	**0.24**	**0.39**	**0.42**	**0.36**	**0.24**	0.15	0.09	-0.02
ISEIND	26.15%	-0.08	**0.29**	**0.44**	**0.48**	**0.41**	**0.28**	0.15	0.08	-0.09
USDTL	8.41%	-0.12	**-0.25**	**-0.37**	**-0.42**	**-0.44**	**-0.36**	-0.33	**-0.27**	-0.06(*)

Notes: Boldfaced coefficients are statistically significant using 0.20 as the threshold value. Coefficients with (*) next to them in the lower part of table reflect a decrease in the coefficient in absolute value compared to the corresponding coefficient in the upper part of the table.

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GLOBAL CAPITAL MARKETS

The global economic growth slowed down in the second quarter accompanied by a decline in trade growth which has affected developments in the major financial markets. Financial markets were volatile, partly reflecting rapidly changing expectations about the duration and depth of the slowdown. In response many countries, especially the United States have eased monetary policy, lowering the target Federal Funds rate three times during the second quarter. The outlook for the GDP in the US in 2001 was lowered slightly to 1.8 %. Japan also revised down the growth forecast to 0.6 % while the forecast for Europe also fell to 2.3 %. In the mature markets equity prices rallied temporarily in April but have since fallen as incoming news on growth prospects deteriorated in all major industrial countries. In currency markets, the US dollar continued to appreciate in nominal effective terms through July, while the euro depreciated by around 7.5 % in nominal terms against the dollar. The yen on the other hand first weakened by 7.5 % (nominal) then picked up for a significant period before returning to the weakening path. The strong U.S. capital inflow might have reflected investor confidence that the US downturn would be short-lived and that high productivity growth would be sustained.

In emerging markets, financial developments have been influenced by mature markets. Reflecting the performance of US equity markets in the second quarter, there was a rally in April and May giving way to a sell off in June. Mutual fund flows into emerging markets also moved in line with the performance of global equity markets but remained volatile due to uncertainty about the timing of a turnaround in global activity. In the emerging debt markets spreads tightened during April and May. Emerging market debt posted a 3.4 % return in the second quarter exceeding that of the US credit markets. Russia was the best performer in the quarter with a return of 16.7 % and Turkey 12.4 % that posted positive returns.

The performances of some developed stock markets with respect to indices indicated that DJI, Nikkei-225 and FTSE-100 decreased by -1.35

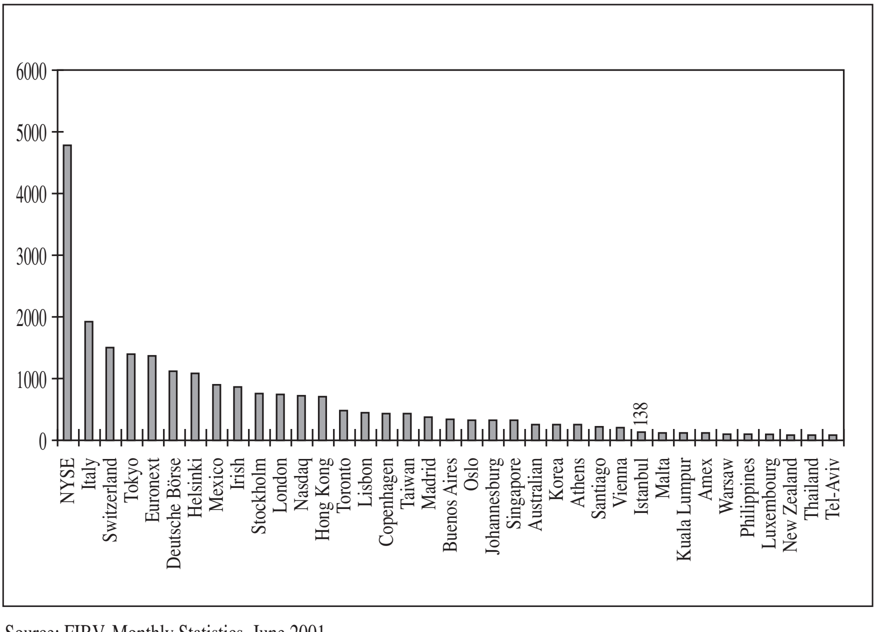
%, -8.62 % and by-5.92 %, respectively, on June 29 as of January 2. When US$ based returns of some emerging markets which were realized between December 31, 2000 and July 4, 2001 are compared, Russia is the best performer with 49.6 %, Mexico follows with 28.8 % and Colombia follows with 20.4 %. In the same period, following Turkey (ISE), with a negative performance by -38.3 %, Brazil, Poland, Egypt, Czech Rep., Israel, India, Singapore and Hungary caused their investors to lose -26.4 %, -22.6 %, -19.4 %, -18.6 % and -18 %, -7.4 %, -16.4 % and -15.8 %, respectively. The performances of emerging markets with respect to P/E ratios as of end-June 2001 indicated that the highest rates were obtained in Malaysia (43.7), Philippines (40.1), Turkey (27.1), Korea (23) and Argentina (23) and the lowest rates in Thailand (-50.5), Indonesia (-8.1), Czech Rep.(6.4), South Africa (9.6) and Brazil (9.9).

Market Capitalization (USD Million, 1986-2000)

	Global	Developed Markets	Emerging Markets	ISE
1986	6,514,199	6,275,582	238,617	938
1987	7,830,778	7,511,072	319,706	3,125
1988	9,728,493	9,245,358	483,135	1,128
1989	11,712,673	10,967,395	745,278	6,756
1990	9,398,391	8,784,770	613,621	18,737
1991	11,342,089	10,434,218	907,871	15,564
1992	10,923,343	9,923,024	1,000,319	9,922
1993	14,016,023	12,327,242	1,688,781	37,824
1994	15,124,051	13,210,778	1,913,273	21,785
1995	17,788,071	15,859,021	1,929,050	20,782
1996	20,412,135	17,982,088	2,272,184	30,797
1997	23,087,006	20,923,911	2,163,095	61,348
1998	26,964,463	25,065,373	1,899,090	33,473
1999	36,030,810	32,956,939	3,073,871	112,276
2000	32,260,433	29,520,707	2,691,452	69,659

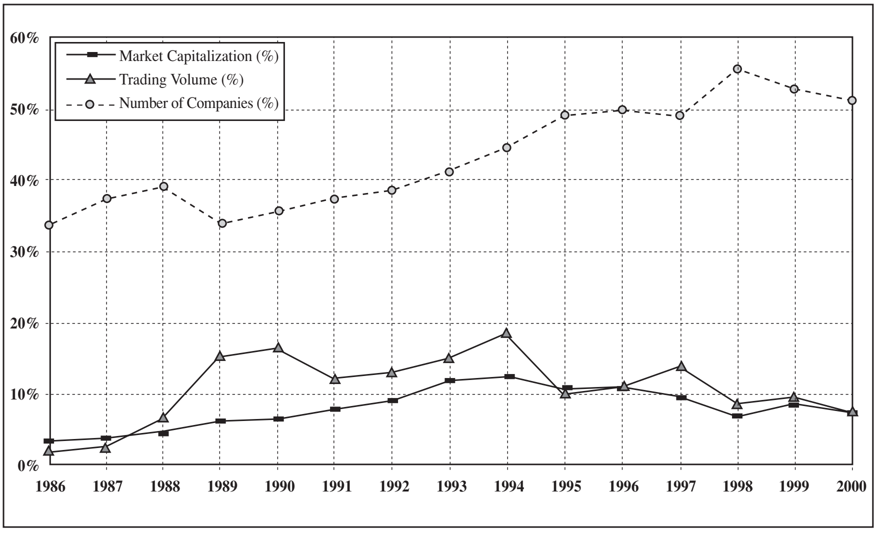
Source: IFC Factbook 2001.

Comparison of Average Market Capitalization (USD Million, June 2001)



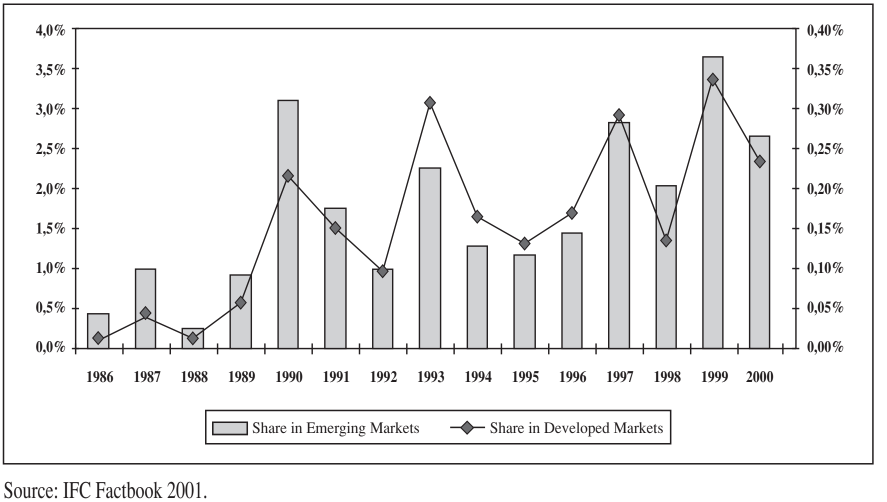
Source: FIBV, Monthly Statistics, June 2001.

Worldwide Share of Emerging Capital Markets (1986-2000)



Source : IFC Factbook, 2001.

Share of ISE's Market Capitalization in World Markets (1986-2000)



Source: IFC Factbook 2001.

Main Indicators of Capital Markets (June 2001)

Market	Turnover Velocity	Market	Value of Share Trading (mill. USD $) Up to Year Total (2001/1-2001/6	Market	Market Cap. of Shares of Domestic Companies (millions USD $)
NASDAQ	396.0%	NASDAQ	6,717,466	NYSE	11,586,436.0
Korea	216.4%	NYSE	5,622,685	NASDAQ	3,210,395.7
Taiwan	205.2%	London	2,505,772	Tokyo	2,984,567.7
Madrid	184.2%	Euronext	1,723,265	London	2,191,614.4
Euronext	175.7%	Tokyo	881,308	Euronext	1,890,172.7
Istanbul	173.9%	Deutsche Börse	826,799	Deutsche Börse	1,110,913.4
Deutsche Börse	119.4%	Amex	477,637	Toronto	668,578.6
Stockholm	110.4%	Chicago	434,260	Switzerland	631,846.3
Italy	109.2%	Madrid	424,859	Hong Kong	579,521.1
Oslo	94.1%	Italy	380,369	Italy	575,658.8
NYSE	87.2%	Switzerland	304,003	Madrid	470,503.9
Copenhagen	86.8%	Taiwan	297,951	Australian	379,175.3
Switzerland	85.9%	Toronto	258,880	Stockholm	241,083.3
Helsinki	83.5%	Stockholm	220,097	Taiwan	231,053.0
London	73.9%	Bermuda	209,566	Johannesburg	194,689.6
Toronto	71.2%	Korea	190,325	Brazil	194,074.4
Thailand	70.7%	Hong Kong	137,882	Korea	182,524.2
Lisbon	61.1%	Australian	119,202	Helsinki	171,491.8
Australian	59.4%	Bilbao	112,672	Mexico	154,905.5
Bilbao	56.2%	Helsinki	101,572	Singapore	127,283.4
Singapore	55.4%	Osaka	96,285	Kuala Lumpur	103,317.0
Warsaw	55.2%	Copenhagen	48,216	Copenhagen	97,583.8
Tokyo	55.0%	Istanbul	46,453	CDNX	91,074.1
New Zealand	53.8%	Singapore	38,008	Irish	81,396.6
Hong Kong	47.0%	Johannesburg	37,997	Athens	80,880.8
Athens	45.7%	Sao Paulo	37,275	Oslo	73,904.8
Sao Paulo	35.4%	Oslo	35,430	Amex	73,456.9
Jakarta	32.6%	Mexico	26,371	Santiago	59,481.5
Irish	31.1%	Barcelona	21,539	Tel-Aviv	57,503.8
Vienna	31.0%	Athens	19,741	Lisbon	45,336.2
Johannesburg	30.9%	Valencia	19,129	Istanbul	43,045.8
Mexico	30.0%	Thailand	15,859	Buenos Aires	41,191.8
Tel-Aviv	29.6%	Lisbon	15,346	Thailand	34,262.7
Ljubljana	24.5%	Irish	11,346	Luxembourg	26,680.3
Philippines	16.1%	Tel-Aviv	7,689	Warsaw	25,564.5
Tehran	15.4%	Kuala Lumpur	7,182	Vienna	24,692.3
Buenos Aires	15.3%	New Zealand	5,975	Philippines	24,562.6
Kuala Lumpur	15.1%	Warsaw	4,994	Jakarta	23,383.5
Lima	13.6%	Jakarta	4,289	New Zealand	17,630.3
Barcelona	10.4%	Buenos Aires	4,217	Lima	10,211.0
Bermuda	10.4%	Vienna	4,184	Tehran	7,009.5
Valencia	9.7%	Santiago	2,199	Ljubljana	2,854.6
Osaka	9.0%	Philippines	1,903	Bermuda	2,641.9
Santiago	8.9%	CDNX	1,600	Malta	1,558.9
Colombo	7.6%	Tehran	568	Colombo	939.0

Source: FIBV. Monthly Statistics. June 2001.

Trading Volume (USD millions, 1986-2000)

	Global	Developed	Emerging	ISE	Emerging/ Global (%)	ISE/ Emerging (%)
1986	3,573,570	3,490,718	82,852	13	2.32	0.02
1987	5,846,864	5,682,143	164,721	118	2.82	0.07
1988	5,997,321	5,588,694	408,627	115	6.81	0.03
1989	7,467,997	6,298,778	1,169,219	773	15.66	0.07
1990	5,514,706	4,614,786	899,920	5,854	16.32	0.65
1991	5,019,596	4,403,631	615,965	8,502	12.27	1.38
1992	4,782,850	4,151,662	631,188	8,567	13.20	1.36
1993	7,194,675	6,090,929	1,103,746	21,770	15.34	1.97
1994	8,821,845	7,156,704	1,665,141	23,203	18.88	1.39
1995	10,218,748	9,176,451	1,042,297	52,357	10.20	5.02
1996	13,616,070	12,105,541	1,510,529	37,737	11.09	2.50
1997	19,484,814	16,818,167	2,666,647	59,105	13.69	2.18
1998	22,874,320	20,917,462	1,909,510	68,646	8.55	3.60
1999	31,021,065	28,154,198	2,866,867	81,277	9.24	2.86
2000	47,869,886	43,817,893	4,051,905	179,209	8.46	4.42

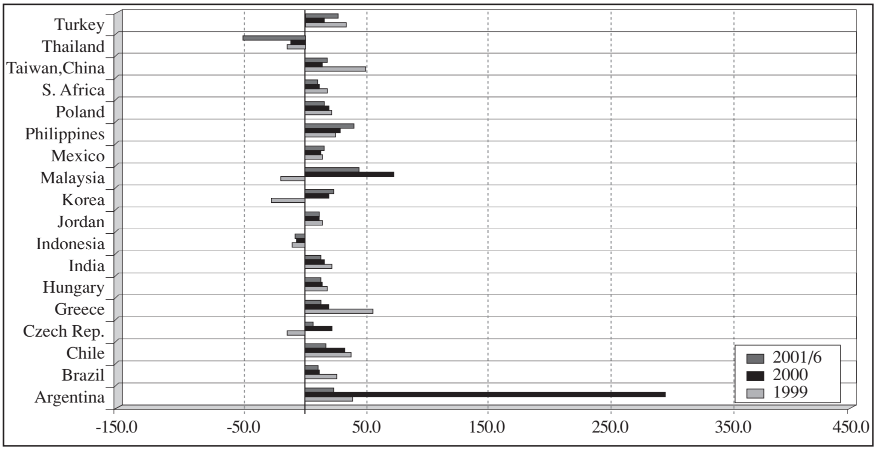
Source: IFC Factbook. 2001.

Number of Trading Companies (1986-2000)

	Global	Developed	Emerging	ISE	Emerging/ Global (%)	ISE/ Emerging (%)
1986	28,173	18,555	9,618	80	34.14	0.83
1987	29,278	18,265	11,013	82	37.62	0.74
1988	29,270	17,805	11,465	79	39.17	0.69
1989	25,925	17,216	8,709	76	33.59	0.87
1990	25,424	16,323	9,101	110	35.80	1.21
1991	26,093	16,239	9,854	134	37.76	1.36
1992	27,706	16,976	10,730	145	38.73	1.35
1993	28,895	17,012	11,883	160	41.12	1.35
1994	33,473	18,505	14,968	176	44.72	1.18
1995	36,602	18,648	17,954	205	49.05	1.14
1996	40,191	20,242	19,949	228	49.64	1.14
1997	40,880	20,805	20,075	258	49.11	1.29
1998	47,465	21,111	26,354	277	55.52	1.05
1999	48,557	22,277	26,280	285	54.12	1.08
2000	49,933	23,996	25,937	315	51.94	1.21

Source: IFC Factbook 2001.

Comparison of P/E Ratios Performances (1999 - 2001/6)



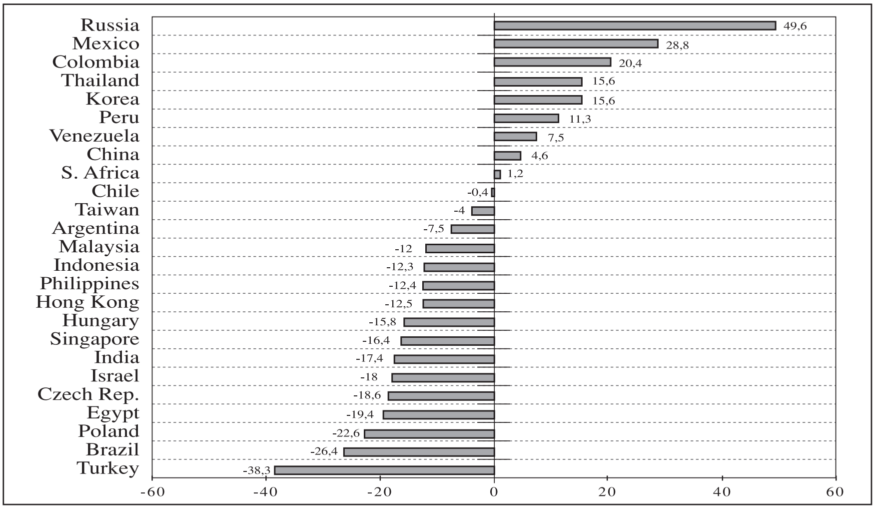
Source: IFC Factbook 2001. IFC, Monthly Review, June 2001.

Price-Earnings Ratios in Emerging Markets (1993-2001/6)

	1993	1994	1995	1996	1997	1998	1999	2000	2001/6
Argentina	41.9	17.7	15.0	38.2	17.1	13.4	39.0	293.3	23.0
Brazil	12.6	13.1	36.3	14.5	15.4	7.0	25.1	11.7	9.9
Chile	20.0	21.4	17.1	27.8	15.9	15.1	37.7	31.8	16.8
Czech Rep.	18.8	16.3	11.2	17.6	8.8	-11.3	-14.8	21.0	6.4
Greece	10.2	10.4	10.5	10.5	13.1	33.7	55.6	19.2	13.1
Hungary	52.4	-55.3	12.0	17.5	25.2	17.0	18.2	14.3	12.1
India	39.7	26.7	14.2	12.3	16.8	13.5	22.0	14.8	12.1
Indonesia	28.9	20.2	19.8	21.6	11.2	-106.2	-10.5	-6.5	-8.1
Jordan	17.9	20.8	18.2	16.9	12.8	15.9	13.6	10.7	11.9
Korea	25.1	34.5	19.8	11.7	11.6	-47.1	-27.7	19.3	23.0
Malaysia	43.5	29.0	25.1	27.1	13.5	21.1	-19.1	71.7	43.7
Mexico	19.4	17.1	28.4	16.8	22.2	23.9	14.1	12.5	15.4
Philippines	38.8	30.8	19.0	20.0	12.5	15.0	24.0	28.2	40.1
Poland	31.5	12.9	7.0	14.3	10.3	10.7	22.0	19.4	15.7
S.Africa	17.3	21.3	18.8	16.3	12.1	10.1	17.4	10.7	9.6
Taiwan, China	34.7	36.8	21.4	28.2	32.4	21.7	49.2	13.7	17.6
Thailand	27.5	21.2	21.7	13.1	4.8	-3.7	-14.5	-12.4	-50.5
Turkey	36.3	31.0	8.4	10.7	18.9	7.8	33.8	15.2	27.1

Source: IFC Factbook. 1999; IFC. Monthly Review. June 2001.
Note: Figures are taken from IFC Investable Index Profile.

Comparison of Market Returns In USD (31/12/2000 - 4/7/2001)



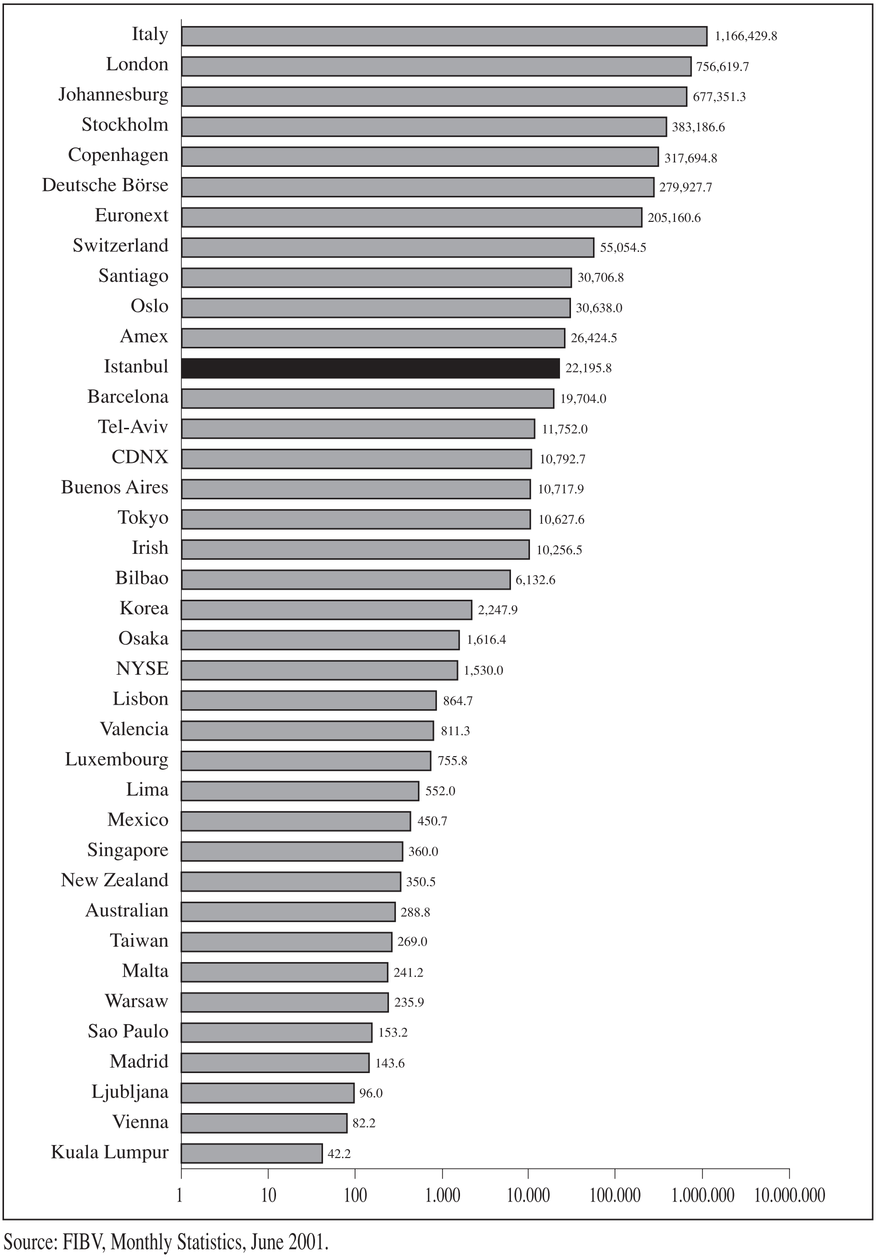
Source: The Economist, July 7th 2001.

Market Vaule/Book Vaule Ratios (1993-2001/6)

	1993	1994	1995	1996	1997	1998	1999	2000	2001/6
Argentina	1.9	1.4	1.3	1.6	1.8	1.3	1.5	1.0	0.9
Brazil	0.5	0.6	0.5	0.7	1.1	0.6	1.6	1.4	1.4
Chile	2.1	2.5	2.1	1.6	1.6	1.1	1.8	1.5	1.6
Czech Rep.	1.3	1.0	0.9	0.9	0.8	0.7	1.2	1.2	0.9
Greece	1.9	1.9	1.8	2.0	2.9	4.9	9.4	4.0	2.3
Hungary	1.6	1.7	1.2	2.0	3.7	3.2	3.6	2.5	1.7
India	4.9	4.2	2.3	2.1	2.7	1.9	3.1	2.5	2.0
Indonesia	3.1	2.4	2.3	2.7	1.5	1.6	2.9	1.6	2.0
Jordan	2.0	1.7	1.9	1.7	1.6	1.8	1.5	1.3	1.3
Korea	1.4	1.6	1.3	0.8	0.6	0.9	2.0	0.8	1.0
Malaysia	5.4	3.8	3.3	3.8	1.8	1.3	1.9	1.5	0.8
Mexico	2.6	2.2	1.7	1.7	2.5	1.4	2.2	1.7	2.2
Philippines	5.2	4.5	3.2	3.1	1.7	1.3	1.5	1.2	1.3
Poland	5.7	2.3	1.3	2.6	1.6	1.5	2.0	2.2	1.8
S.Africa	1.8	2.6	2.5	2.3	1.9	1.5	2.7	2.1	2.2
Taiwan, China	3.9	4.4	2.7	3.3	3.8	2.6	3.3	1.7	1.7
Thailand	4.7	3.7	3.3	1.8	0.8	1.2	2.6	1.6	2.0
Turkey	7.2	6.3	2.7	4.0	9.2	2.7	8.8	3.1	3.9

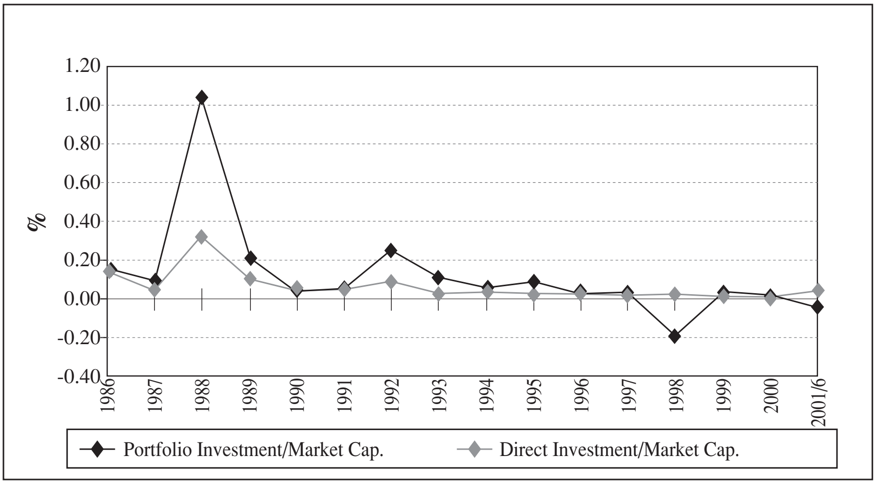
Source: : IFC Factbook. 1996-2001; IFC Monthly Review, June 2001.

Value of Bond Trading (Million USD. January 2001-June 2001)



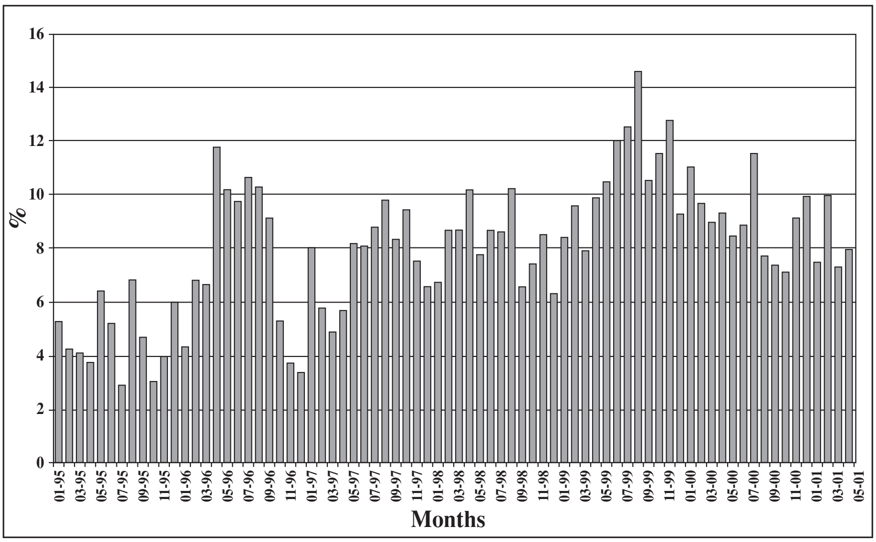
Source: FIBV, Monthly Statistics, June 2001.

Foreign Investments as a Percentage of Market Capitalization in Turkey (1986-2001/6)



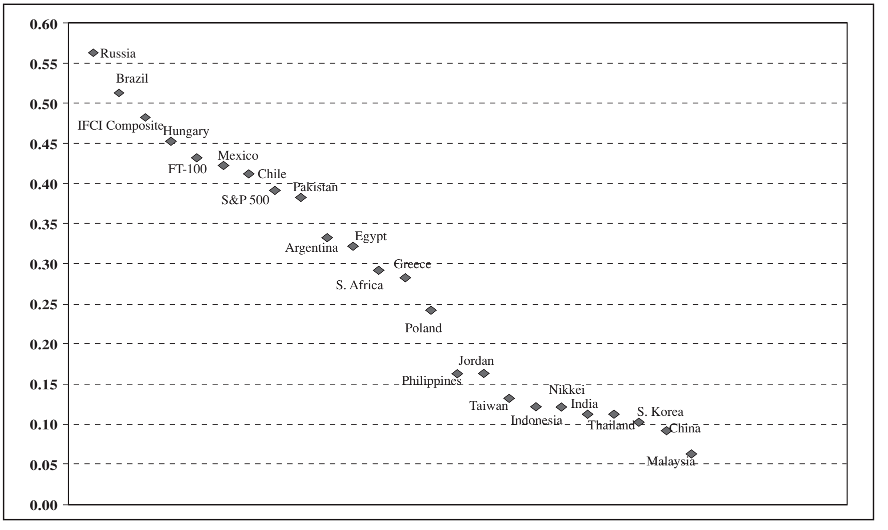
Source: ISE Data. CBTR Databank.

Foreigners' Share in the Trading Volume of the ISE (Jan. 1995-June 2000)



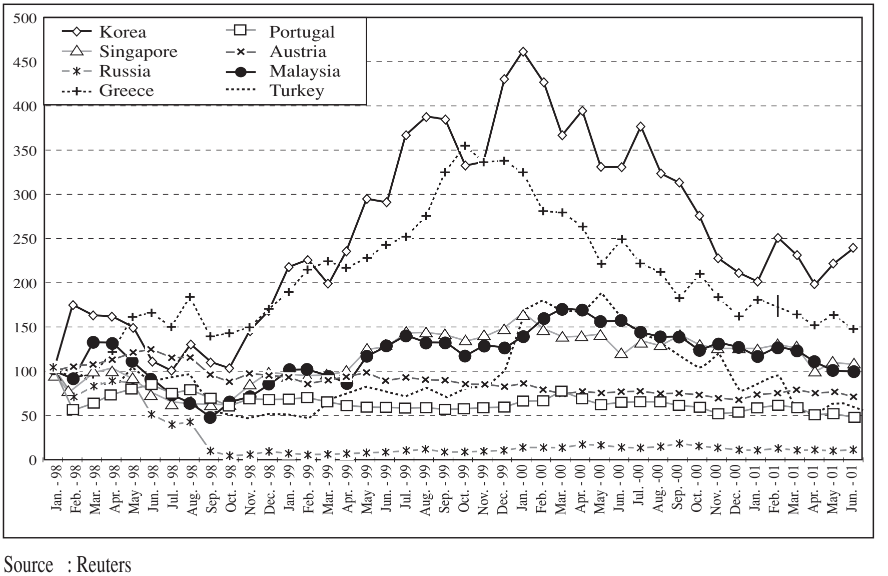
Source: ISE Data.

Price Correlations of the ISE (June 1997-June 2001)



Source : IFC Monthly Review, June 2001.
Notes : The correlation coefficient is between -1 and +1. If it is zero. for the given period. it is implied that there is no relation between two serious of returns. For monthly return index correlations (IFCI) see. IFC. Monthly Review. Oct. 1999.

Comparison of Market Indices (31 Dec. 1997 =100)



Source : Reuters
Note : Comparisons are in US$.

The ISE Review Volume: 5 No: 18 April/May/June 2001
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ISE
Market Indicators

STOCK MARKET											
		Total Value				Market Value		Dividend Yield	P/E Ratios		
	Number of Companies	Total		Daily Average							
		(TL Billion)	(US$ Million)	(TL Billion)	(US$ Million)	(TL Billion)	(US$ Million)	(%)	TL(1)	TL(2)	US $
1986	80	9	13	—	—	709	938	9.15	5.07	—	—
1987	82	105	118	—	—	3,182	3,125	2.82	15.86	—	—
1988	79	149	115	1	—	2,048	1,128	10.48	4.97	—	—
1989	76	1,736	773	7	3	15,553	6,756	3.44	15.74	—	—
1990	110	15,313	5,854	62	24	55,238	18,737	2.62	23.97	—	—
1991	134	35,487	8,502	144	34	78,907	15,564	3.95	15.88	—	—
1992	145	56,339	8,567	224	34	84,809	9,922	6.43	11.39	—	—
1993	160	255,222	21,770	1,037	88	546,316	37,824	1.65	25.75	20.72	14.86
1994	176	650,864	23,203	2,573	92	836,118	21,785	2.78	24.83	16.70	10.97
1995	205	2,374,055	52,357	9,458	209	1,264,998	20,782	3.56	9.23	7.67	5.48
1996	228	3,031,185	37,737	12,272	153	3,275,038	30,797	2.87	12.15	10.86	7.72
1997	258	9,048,721	58,104	35,908	231	12,654,308	61,879	1.56	24.39	19.45	13.28
1998	277	18,029,967	70,396	72,701	284	10,611,820	33,975	3.37	8.84	8.11	6.36
1999	285	36,877,335	84,034	156,260	356	61,137,073	114,271	0.72	37.52	34.08	24.95
2000	315	111,165,396	181,934	451,892	740	46,692,373	69,507	1.29	16.82	16.11	14.05
2001	312	46,477,074	48,454	377,862	394	54,022,417	43,152	1.21	22.25	36.73	17.74
2001/Ç1	315	18,110,652	24,208	306,960	410	40,039,488	39,260	1.46	17.07	17.23	10.42
2001/Ç2	312	28,366,421	24,246	443,225	379	54,022,417	43,152	1.21	22.25	36.73	17.74

Q: Quarter
Note:
- Between 1986-1992, the price earnings ratios were calculated on the basis of the companies' previous year-end net profits. As from 1993,
  TL(1) = Total market capitalization / Sum of last two six-month profits
  TL(2) = Total market capitalization / Sum of last four three-month profits.
  US$ = US$ based total market capitilization / Sum of last four US$ based three-month profits.

Closing Values of the ISE Price Indices

	TL Based				
	NATIONAL-100 (Jan. 1986=1)	NATIONAL-INDUSTRIALS (Dec. 31, 90=33)	NATIONAL-SERVICES (Dec. 27, 96=1046)	NATIONAL-FINANCIALS (Dec. 31, 90=33)	NATIONAL-TECHNOLOGY (June, 30,2000=14.466,12)
1986	1.71	—	—	—	—
1987	6.73	—	—	—	—
1988	3.74	—	—	—	—
1989	22.18	—	—	—	—
1990	32.56	32.56	—	32.56	—
1991	43.69	49.63	—	33.55	—
1992	40.04	49.15	—	24.34	—
1993	206.83	222.88	—	191.90	—
1994	272.57	304.74	—	229.64	—
1995	400.25	462.47	—	300.04	—
1996	975.89	1,045.91	1,046.00	914.47	—
1997	3,451.—	2,660.—	3,593.—	4,522.—	—
1998	2,597.91	1,943.67	3,697.10	3,269.58	—
1999	15,208.78	9,945.75	13,194.40	21,180.77	—
2000	9,437.21	6,954.99	7,224.01	12,837.92	10,586.58
2001	11,204.24	8,702.96	6,524.80	16,045.49	7,914.39
2001/Q1	8,022.72	6,395.44	5,369.60	10,827.58	7,633.62
2001/Q2	11,204.24	8,702.96	6,524.80	16,045.49	7,914.39

	US $ Based					EURO Based
	NATIONAL-100 (Jan. 1986=100)	NATIONAL-INDUSTRIALS (Dec. 31, 90=643)	NATIONAL-SERVICES (Dec. 27, 96=572)	NATIONAL-FINANCIALS (Dec. 31, 90=643)	NATIONAL-TECHNOLOGY (Jun. 30, 00=1.360.92)	NATIONAL-100 (Dec.31,98=484)
1986	131.53	—	—	—	—	—
1987	384.57	—	—	—	—	—
1988	119.82	—	—	—	—	—
1989	560.57	—	—	—	—	—
1990	642.63	642.63	—	642.63	—	—
1991	501.50	569.63	—	385.14	—	—
1992	272.61	334.59	—	165.68	—	—
1993	833.28	897.96	—	773.13	—	—
1994	413.27	462.03	—	348.18	—	—
1995	382.62	442.11	—	286.83	—	—
1996	534.01	572.33	572.00	500.40	—	—
1997	982.—	757.—	1,022.—	1,287.—	—	—
1998	484.01	362.12	688.79	609.14	—	484.01
1999	1,654.17	1,081.74	1,435.08	2,303.71	—	1,912.46
2000	817.49	602.47	625.78	1,112.08	917.06	1,045.57
2001	520.80	404.53	303.29	745.83	367.88	718.60
2001/Q1	457.77	364.91	306.38	617.81	435.56	607.16
2001/Q2	520.80	404.53	303.29	745.83	367.88	718.60

Q : Quarter

* The second quarter figures are as of June 29, 2001.

BONDS AND BILLS MARKET

Traded Value

Outright Purchases and Sales Market

	Total		Daily Average	
	(TL Billion)	(US$ Million)	(TL Billion)	(US$ Million)
1991	1,476	312	11	2
1992	17,977	2,406	72	10
1993	122,858	10,728	499	44
1994	269,992	8,832	1,067	35
1995	739,942	16,509	2,936	66
1996	2,710,973	32,737	10,758	130
1997	5,503,632	35,472	21,840	141
1998	17,995,993	68,399	71,984	274
1999	35,430,078	83,842	142,863	338
2000	166,336,480	262,941	662,695	1,048
2001	19,932,249	23,728	160,744	191
2001/Ç1	11,798,611	16,825	196,644	280
2001/Ç2	8,133,638	6,902	127,088	108

Repo-Reverse Repo Market

	Total		Daily Average	
	(TL Billion)	(US$ Million)	(TL Billion)	(US$ Million)
1993	59,009	4,794	276	22
1994	756,683	23,704	2,991	94
1995	5,781,776	123,254	22,944	489
1996	18,340,459	221,405	72,780	879
1997	58,192,071	374,384	230,921	1,486
1998	97,278,476	372,201	389,114	1,489
1999	250,723,656	589,267	1,010,982	2,376
2000	554,121,078	886,732	2,207,654	3,533
2001	376,278,043	405,024	3,034,500	3,266
2001/Q1	191,773,165	249,085	3,196,219	4,151
2001/Q2	184,504,878	155,939	2,882,889	2,437

Q : Quarter

ISE GDS Price Indices (December 25-29, 1995=100)

	TL Based			
	30 Days	**91 Days**	**182 Days**	**General**
1996	103.41	110.73	121.71	110.52
1997	102.68	108.76	118.48	110.77
1998	103.57	110.54	119.64	110.26
1999	107.70	123.26	144.12	125.47
2000	104.84	117.12	140.81	126.95
2001	106.50	119.22	135.40	118.58
2001/Q1	103.38	109.26	115.47	108.00
2001/Q2	106.50	119.22	135.40	118.58

ISE GDS Performance Indices (December 25-29, 1995=100)

	TL Based		
	30 Days	**91 Days**	**182 Days**
1996	222.52	240.92	262.20
1997	441.25	474.75	525.17
1998	812.81	897.19	983.16
1999	1,372.71	1,576.80	1,928.63
2000	1,835.26	2,020.94	2,538.65
2001	2,428.18	2,724.69	3,114.67
2001/Q1	2,160.79	2,270.15	2,595.08
2001/Q2	2,428.18	2,724.69	3,114.67
	US $ Based		
1996	122.84	132.99	144.74
1997	127.67	137.36	151.95
1998	153.97	169.96	186.24
1999	151.02	173.47	212.18
2000	148.86	169.79	213.28
2001	114.76	128.77	147.21
2001/Q1	125.36	131.71	150.56
2001/Q2	114.76	128.77	147.21

Q : Quarter

* The second quarter figures are as of June 29, 2001.

ISE GDS Price Indices (January 02, 2001=100)

	TL Based				
	6 Months (182 Days)	9 Months (273 Days)	12 Months (365 Days)	15 Months (456 Days)	General
2001	98.02	92.50	85.72	78.63	95.78
2001/Q1	88.08	82.14	76.35	70.90	81.62
2001/Q2	98.02	92.50	85.72	78.63	95.78

ISE GDS Performance Indices (January 02, 2001=100)

	TL Based			
	6 Months (182 Days)	9 Months (273 Days)	12 Months (365 Days)	15 Months (456 Days)
2001	131.88	126.18	117.03	107.17
2001/Q1	106.09	97.04	88.65	81.53
2001/Q2	131.88	126.18	117.03	107.17
	US $ Based			
2001	6.23	5.96	5.53	5.07
2001/Q1	6.15	5.63	5.14	4.73
2001/Q2	6.23	5.96	5.53	5.07

Q : Quarter

* The second quarter figures are as of June 29, 2001.

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Book Reviews

"*Stock Markets, Speculative Bubbles and Economic Growth: New Dimensions in the Co-Evolution of Real and Financial Markets*", Mathias Binswanger, Edward Elgar Publishing. 2001, Cheltenham., pp.vii-359.

The author explains the importance of the relation between the speculative bubbles and the economic development, thereby concluding that those bubbles cannot easily be corrected by policy measures. The major argument presented in this book is that several developments during 1980's, such as financial innovations and institutional arrangements, enabled the emergence of sustainable speculative bubbles on the stock market in the USA during the 1980's.

This book attempts to show that recent developments in financial markets , in the US and probably also in other industrial countries with well-developed financial sectors, caused a relaxation of the aggregate finance where it contributed to the emergence of speculative bubles. But speculative bubles themselves may be understood as a response to operative real and demand constraints. Consequently, the co-evolution between the real and fnancial sectors in the economy became more complex as speculative bubles also influence the level of real economic activity.

The book includes four main parts covering ten chapters following the introduction.

In the introduction part stresses that the whole development cannot be explained in a coherent analytical framework within the realm of one theory and that the help is needed of a variety of economic theories for the different aspects of economic development. This book gives the constraints in the economic development which can be distinguished in three main categories as the finance constraint, the real constraint and the demand constraint.

The alterations made in the banking sector relaxed the finance constraint. Banks' credits, financial innovations and deregulation further facilitated the finance constraint. So the role of the financial sector in economic development is stressed. Real constraints arise when there is a scarcity of profitable investment opportunities. Sometimes existing production possibilities are not sufficient to create enough prospects for profits. Another real constraint is due to the fact that the removal of the finance constraint leading to the increase in financial capital, is much faster than innovations and profit opportunities in production, which are seen as a cure for real constraint. However, without an increase in production, speculative bubbles contribute to a relaxation of the real constraint as they increase the aggregate profits.

The third category, the demand constraint arises from the discrepancy between expected and actual developments of the economy. Firms never accurately predict how much consumers and investors will spend for their products. There is always the possibility that the demand is much lower than the expected level. To hinder the overcapacities, firms reduce investment in real capital. So the demand aspect may constrain economic growth.

The first part of the book covers the chapters from 1 to 3 and concentrates on the role of finance and money in the economic process. Shortly, chapter 1 argues that financial sector services promotes investment which in turn, leads to growth. In this chapter, Keynesian and non-Keynesian approaches over this issue is also given. Chapter 2 describes the money creation in contemporary economies while chapter 3 focuses on the issue that relaxation of credit constraint accelerates investment in financial assets which increases money flows. In chapter 3, these money flows are interpreted as financial hoarding.

Second part, covering chapters 4 to 6, examines speculative bubbles on the stock market and describes the conditions under which bubbles may occur. Chapter 5 and 6 states that institutional changes and financial innovations promote confidence and speculative bubbles may only be sustainable, if there is a high level of confidence into the functioning of financial sector.

The third part, entailing chapters 7 to 9, puts forth empirical evidence concerning the recent development in the U.S. Chapter 7 mainly on the stock markets, links the development of stock prices to the 1980's merger wave and emphasizes information asymmetries on the stock market. Chapter 8 and 9 gives empirical evidence showing that investment in the real sector declined due to low profits in the 1980's and for the emergence of speculative bubbles in the US economy.

The fourth part (chapter 10) concludes by linking different parts of the book together and provide more empirical evidence.

ISE PUBLICATIONS

I- PERIODICALS	ISSN/ISBN	DATE
Weekly Bulletin	ISSN 1300-9311	
Monthly Bulletin (Turkish)	ISSN 1300-9303	
Monthly Bulletin (English)	ISSN 1300-9834	
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