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The relationship between stock and real estate prices in Turkey: Evidence around the global financial crisis

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ABSTRACT

Research on the relationship between stock and real estate prices focuses on two transmission mechanisms, namely the wealth and credit-price effects. This paper uses the 2007 global financial crisis as a natural experiment and examines whether the relationship between real estate prices and stock prices has changed after the outbreak of the crisis by using data from the Turkish market. The results based on a threshold cointegration framework indicate that while both effects exist during the pre-crisis period, only a credit-price effect is observed during the crisis period. Moreover, the findings are sensitive to whether or not one allows for asymmetric error correction.

1. Introduction

Research on the impact of wealth on consumption has attracted interest on the relationship between stock and real estate prices. In the resulting literature, two mechanisms have been proposed that may lead to a causal relationship between real estate prices and stock prices. The first mechanism, known as the wealth effect, argues that as the stock market rises, investors with unanticipated increases in wealth will push their demand for real estate up. Hence, the stock market will lead the real estate market. The second mechanism, the so called credit-price effect, emphasizes that real estate serves as collateral to especially credit-constrained firms. An increase in real estate prices would improve the balance-sheet position of these firms and decrease their costs of borrowing. This will lead to a higher level of investment activity by firms accompanied by a rise in their stock prices. Based on this reasoning, the credit-price effect predicts that the real estate market will lead the stock market. Overall, the majority of the empirical research provides supporting evidence for the wealth effect.[1]

Stock and real estate markets are, however, also affected by economic conditions. For example, the 2007 global financial crisis that began in the U.S. had a considerable negative effect on both stock and real estate prices in many countries. As Lin and Treichel (2012) describes “The ensuing financial sector crisis quickly led to a significant decline in credit to the private sector as well as to a sharp rise in interest rates. The resulting collapse in U.S. financial institutions led to a collapse of equity markets, and of international trade and industrial production and spread to other advanced economies as well as to emerging markets and developing countries. Real growth around the world declined sharply below projections and advanced economies, including the U.S., entered into a recession”. Adair et al. (2009) argues that globally \$7 trillion has been wiped off the stock markets over the course of 2008.[2] Another consequence of the crisis is the collapse of the housing market in many countries.[3]

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[1] Green (2002), Kapopoulos and Siokis (2005), Chen (2001), Sutton (2002), Kakes and Van Den End (2004), Ibrahim (2010) provide evidence consistent with the existence of a wealth effect. Lean and Smyth (2012), Liu and Su (2010) and Su (2011) report the existence of both effects. Only, Sim and Chang (2006) find the existence of credit effect.

[2] During 2008, New York's S&P 500, Japan's Nikkei 225 and UK's FTSE 100 indices fell by 38.5%, 42% and 31.3%, respectively.

[3] In the US, The Case-Shiller Index shows a 25% drop over the two year period 2007–2008 (Barker, 2009). According to Nationwide Building Society, for example, the average house price in the UK fell by 14.7% over the course of 2008 (Adair et al., 2009).

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During the first quarter of the 2008, the Turkish stock market, with a decline of 36.62 percent, showed the most drastic reaction to the financial crisis among the countries included in the S&P/Citigroup BMI Global Index (Standard and Poor's, 2008). Moreover, the REIDIN Turkey Residential Property Price Indices show a continual fall in house prices over the period March 2008 to March 2009 (ReidinTurkey, 2010). The overall performance of the economy was unsatisfactory during these two year period, with GDP growth rates of 0.7 percent and 4.7 percent, respectively for 2008 and 2009 (Turkey's Statistical Yearbook, 2009). In spite of the dramatic initial reaction, the Turkish economy has shown recovery by the end of 2010. Coskun (2011) argues "Turkey has faced limited negative impacts from the global financial crisis. The lack of securitization/structured product markets and also inefficient housing credit market may have seemed good news for Turkey during the financial turmoil."

The purpose of this paper is to use this period of crisis as a natural experiment and examine whether the relationship between real estate and stock prices in Turkey has changed following the crisis. While this relationship has been extensively researched in a number of international markets, no attempt has been made to assess Turkey. We examine this issue by using daily Real Estate Investment Trusts (REIT) index, stock market index and interest rate data within the framework of a vector error correction model. Since, as argued in Gonzalo and Pitarakis (2006), omitting the presence of nonlinear components, like threshold effects in long-run equilibrium, can lead to misinterpretations of equilibrium relationships, we employ threshold cointegration where adjustments to a long-run equilibrium only takes place when deviations become large and exceed the threshold.

The results indicate that the nature of the long-run relation has changed after the outbreak of the crisis. When we compare pre-crisis and crisis periods, it is noteworthy that while both credit-price and wealth effects exist during the pre-crisis period, only a credit-price effect is observed during the crisis period. Moreover, the findings are sensitive to whether or not one allows for asymmetric error correction.

The remainder of the paper is organized as follows. The next section gives a brief review of the literature. The third section presents the data and methodology used in the study. The fourth section reports and discusses the empirical results. The last section provides the concluding remarks.

2. Literature review

Early research on the relationship between real estate and stock prices has examined correlations between the returns on these two investment alternatives using U.S. or U.K. data.[4] Most of these articles report that real estate and stock returns are negatively correlated. However, the evidence in these studies cannot be used to distinguish between the credit-price or wealth effect, since they do not test for the direction of causality.

The next wave of studies applies the concept of Granger causality, vector autoregression (VAR) modeling and cointegration techniques to examine the causal interactions between real estate and stock prices. Green (2002) and Kapopoulos and Siokis (2005) test for Granger causality using a single-equation framework. Chen (2001), Sutton (2002), Kakes and Van Den End (2004) and Sim and Chang (2006) use VAR modeling. Ibrahim (2010) and Lean and Smyth (2012) use the cointegration technique and vector error correction modeling. Overall, only Sim and Chang (2006) and Lean and Smyth (2012) provide supporting evidence for the credit-price effect. Each of the other studies, though, favors the wealth effect.

Both Green (2002) and Kapopoulos and Siokis (2005) use differenced series in a single equation framework. Green (2002) uses data from San Francisco Bay area, which is argued to be a prime candidate for a wealth effect to be large for the following reasons. First, high income households in this region are expected to hold relatively large amounts of stock. Second, compared to workers elsewhere workers in San Francisco Bay area are more likely to be paid in stocks. The paper provides evidence, consistent with this conjecture that stock prices Granger cause real estate prices in this region. Similarly, Kapopoulos and Siokis (2005) reports evidence in favor of the wealth effect hypothesis for Athens real estate prices, but not for other urban real estate prices.

One of the studies that use a VAR framework, Chen (2001), examines the relation in Taiwanese market by including rediscount rates and the total amount of bank loans as control variables. The findings indicate that stock prices Granger cause housing prices, but not vice versa. Moreover, changes in bank loans, but not changes in rediscount rates, are significant in predicting both stock and housing prices. Another study, Sutton (2002), examines the extent to which house price changes in six developed markets, namely the United States, the United Kingdom, Canada, Ireland, the Netherlands and Australia, can be explained by changes in national incomes, interest rates and stock prices. This study also provides evidence in favor of the wealth effect.

A further study that uses VAR modeling, Kakes and Van Den End (2004), examines the Dutch market using real disposable income and the ten-year government bond yield as control variables. The findings show that stock prices and interest rate have explanatory power for future changes in house prices. Moreover, their results indicate the absence of a credit-price effect in the Dutch market. The results also show, consistent with the evidence in Green (2002), that an increase in homeowners' participation in the stock market increases the sensitivity of house prices to stock market. A similar paper, Sim and Chang (2006), examines the Korean market by using the growth rate of GDP and three-year corporate bond yield as control variables. Overall, the results show that house and land prices Granger-cause stock prices in most regional housing and land markets, but there is no converse causation from stock to real estate markets. The findings reveal that the credit-price effect is particularly associated with industrial land markets.

The first of the two papers that use a cointegration framework, Ibrahim (2010), examines the relation between stock and real estate prices in the Thai market. By including real output and consumer price data in the analysis, the paper finds strong evidence in favor of a wealth effect. It also documents that real activity has significant impact on both real estate and stock prices. The second paper, Lean and Smyth (2012), examines Malaysia by employing interbank deposit rates as control variable. The paper uses individual REIT rather than REIT index data. While a wealth effect is found for some REITs, for most of the others there is evidence of feedback effects between real estate and stock markets.

Other recent studies, recognizing that standard cointegration technique fails to capture real world economic phenomena such as the possible impact of market frictions, asymmetric information and transaction costs on the adjustment to the long-run equilibrium, employ threshold cointegration modeling. This method, in which adjustments to a long-run equilibrium only take place when deviations become large and exceed the threshold, has been popular after the seminal paper of Balke and Fomby (1997). As Gonzalo and Pitarakis (2006) points out, omitting the presence of nonlinear components, like threshold effects in long-run equilibrium, can

[4] See, for example, Ibbotson and Siegel (1984), Hartzell (1986) and Eichholtz and Hartzell (1996).

lead to misinterpretations of equilibrium relationships because the cointegrating vector will no longer be consistently estimated.

Using this framework, Liu and Su (2010) examines the bivariate relationship between real estate and stock market indices in China. The paper's findings indicate bidirectional causality. In other words, both credit-price and wealth effects exist in China. In a similar study, Su (2011) uses data from eight Western European countries. The results of Granger causality test differs across the countries examined. While, a credit-price effect is observed in Germany, the Netherlands and the UK, a wealth effect is found in Belgium and Italy. Moreover, both effects exist in France, Spain and Switzerland.

3. Data and methodology

Data used in this study contains daily closing values of REIT index, stock market index and the one-month deposit rate. Using REIT data is a popular alternative for purposes of examining the linkages between stock and real estate investments (Subrahmanyam, 2007). In contrast to unit trusts, REITs are actively traded on stock exchanges. Hoesli and Oikarinen (2012) using data from the U.S., U.K., and Australia find that the long-run REIT market performance is much more closely related to the direct real estate market than to the general stock market.[5]

The data are retrieved from Datastream. The sample period covers 1000 day period around the outbreak of the global financial crisis of 2007. The interval November 14, 2005 to October 14, 2007 is classified as the pre-crisis period, whereas the October 15, 2007 to September 11, 2009 time frame is classified as the crisis period.

Founded in December 1985, Borsa Istanbul (BIST) showed a substantial growth until the global financial crisis. At the beginning of the sample period, total market capitalization of the 300 stocks listed on BIST was $98 billion.[6] As of the end of 2009, 316 listed stocks had a total market capitalization of $236 billion. REIT practices started in Turkey in 1995, while REIT shares have been traded on BIST since 1997. Just like their counterparts in other countries, Turkish REITs are exempt from both corporate and income taxes. However, unlike their counterparts, they are not required to pay out dividends to their shareholders on an annual basis.

At the beginning of the sample period, there were 10 REITs listed on BIST. They made up about 1.21% of the total Turkish stock market capitalization. As of the end of 2009, the number of REITs increased to 14 with a share of 2.08% in total market capitalization. Table 1 gives the same figure for selected European and Asian countries at the end of 2012. The share of REITs in these countries ranges from 0.02% (in South Korea) to 3.77% (in France) and hence is not very different than that in Turkey.

Summary statistics for REIT index, stock market index and interest rate are presented in Table 2. As the table shows, the level of the REIT index falls considerably during the crisis period. On the other hand, stock market index and interest rate show minor decreases.[7] Another notable observation is that the volatility increases for all the variables. The skewness values show that REIT and stock indices have long right tails, while interest rate has a long left tail in both periods. The kurtosis values indicate that all three distributions are flat relative to the normal in both periods. All three distributions are non-normal during both periods.

Fig. 1 shows the monthly average values for the three variables. It is notable that REIT and stock market indices move together in both periods. As the graph shows that both REIT and stock market indices fall considerably over the period from November 2007 to November 2008. In fact, the choice of the starting date of the crisis as October 15, 2007 is based on this observation.[8] To check the robustness of our results to the choice of this date, the analyses will be repeated by choosing October 1, 2007 and November 1, 2007 as the starting date of the crisis.

Fig. 1 also shows that over the sample period there were two striking trends in interest rates. First, the rise that starts around May 2006 and second the fall that starts during the last quarter of 2008. Both of these periods are characterized by capital outflows from developing countries, including Turkey. During these periods there was significant depreciation of the Turkish Lira. It depreciated by around 20 percent during the period May–June 2006 and 30 percent during the last quarter of 2008. The monetary policy decisions of the Central Bank in these two occasions are in contrast. While an unexpected and strong monetary tightening was implemented in the face of capital outflows in the May–June 2006 period, an expansionary monetary policy was adopted following capital outflows in the last quarter of 2008.

Yilmaz (2008) explains the contradiction between these two responses mainly by the significant differences between the two periods in terms of both economic conjuncture and the sources of the shocks. As both domestic demand and external demand were strong in 2006, the exchange rate movements triggered by capital outflows had the potential lead to an increase in expected inflation. However, in 2008, the global recession and weak domestic demand conditions exerted downward pressure on the inflation. The Central Bank reduced short-term interest rates, in order to offset the extra tightening in monetary conditions. [9] All of these suggest that the

Table 1
The market share of REITs in selected European and Asian countries.

	REIT (billion $)	Stock market (billion $)	Share (%)
European countries			
Belgium	7.6	300	2.54
France	68.8	1823	3.77
Germany	1.5	1486	0.10
Greece	0.6	45	1.31
Italy	1.5	480	0.31
Netherlands	9.1	651	1.40
Turkey	7.8	309	2.52
UK	42.5	3019	1.41
Asian countries			
Hong Kong	22.2	1108	2.00
Japan	52.2	3681	1.42
Malaysia	8.0	476	1.68
South Korea	0.2	1180	0.02
Thailand	6.1	383	1.59

Notes: The figures are as of the end of 2012. REIT market capitalization figures are from Ooi and Wong (2013) and Mattarocci (2014), total stock market capitalization figures are from Quandl (http://www.quandl.com).

[5] A House Price Index, which covers the whole country, is constructed by the Central Bank of Turkey for the purpose of monitoring price movements in the Turkish housing market. This index is introduced in January 2010 later than the end of the sample period in this study. An informal comparison (not shown) of the house price index to the REIT index over the period January 2010 to February 2014 indicates that the REIT index is much more volatile than the House Price Index.

[6] REIT and stock market figures in this paragraph are obtained from BIST http://www.borsaistanbul.com/. Interest rate figures are obtained from Central Bank of Turkey, electronic data delivery system http://evds.tcmb.gov.tr/.

[7] The sample period is characterized by a level of interest rates which is dramatically lower than those during the past three decades. After the banking crisis that occurred in 2001, a new economic program set fighting inflation as a main goal of economic policy. As a result average interest rates for one-month bank deposits fall from 65% at the end of 2001 to around 19% at the beginning of 2005.

[8] A large number of papers (see e.g. Naifar, 2011) take August 9, 2007 as the starting date for the global financial crisis, when BNP Paribas terminated withdrawals from three hedge funds citing “a complete evaporation of liquidity”. Nonetheless, the starting date of the crisis may differ across countries.

[9] Real GDP growth rates over 2005–2007 period are 8.4%, 6.9%, and 4.7%, while they are 0.7% and −4.8% in 2008 and 2009, respectively (IMF World Economic Outlook, 2015).

Table 2
Descriptive statistics.

	Pre-crisis			Crisis		
	REIT index	Stock index	Interest rate	Reit index	Stock index	Interest rate
Mean	34,764.91	42,346.61	17.32	21,470.25	37,810.09	14.73
Median	34,419.48	41,974.57	18.35	21,884.75	38,213.75	16.15
Std. Dev.	4035.04	5267.40	2.38	7575.48	9688.86	3.45
Skewness	0.31	0.56	−0.59	0.45	0.22	−0.69
Kurtosis	2.72	2.91	1.84	2.47	2.18	1.98
Jarque–Bera	9.37[a]	26.3[a]	57.22[a]	22.96[a]	18.11[a]	60.9[a]

Notes: For Jarque–Bera test [a] indicates significance at 1% level.

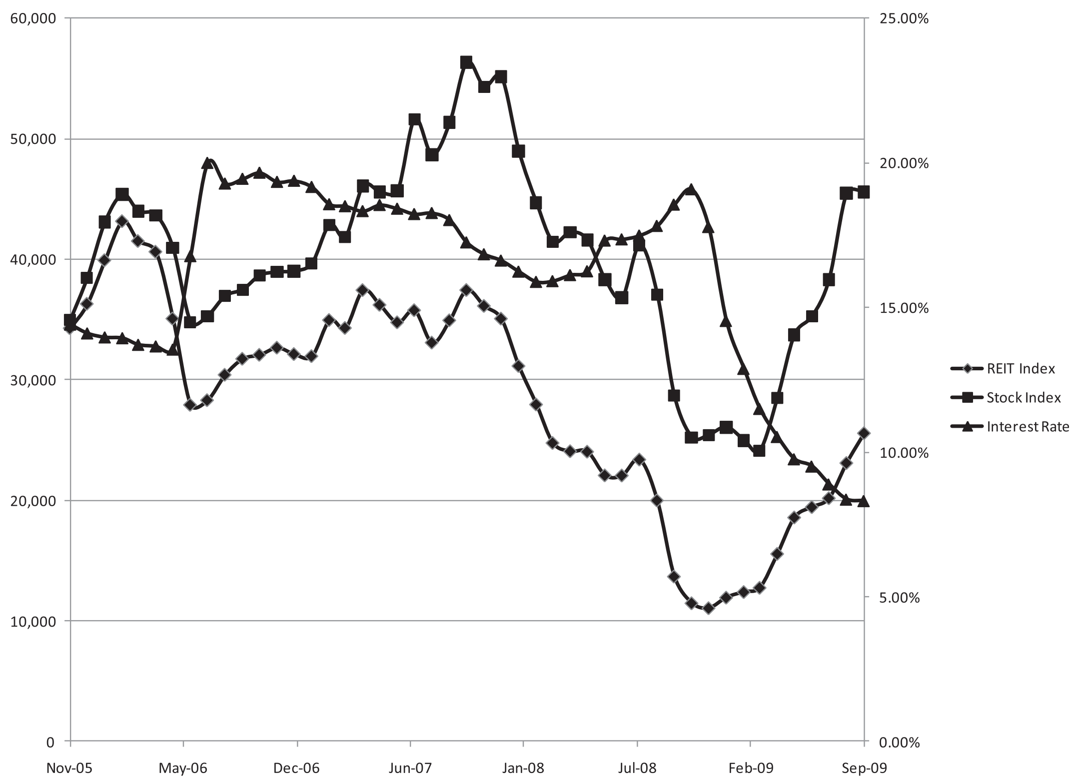

Fig. 1. Monthly average values for REIT index, stock index and interest rate.

economic conjuncture was quite different during the pre crisis and crisis periods.

All data are transformed to natural logarithms before the analysis. Before outlining the methodology, we pretest the variables for unit roots and stationarity by using Augmented Dickey-Fuller (ADF) and Phillips–Perron (PP) tests. The results presented in Table 3 indicate that both tests cannot reject the null of a unit root for all series in level. However, they confirm stationarity when all the series are in the first difference. As a result, we conclude that the series are all I(1), which is the premise of cointegration.

Our methodology starts with employing the VAR-based approach of Johansen (1988) and Johansen and Juselius (1990) to test for cointegration or long-run relationship among the variables. The findings of this method, which assumes linear behavior in the long and short run, will serve as a benchmark for comparison. As it was mentioned earlier in the text, omitting the presence of nonlinear components, like threshold effects in long-run equilibrium may lead to inconsistency of the estimated equilibrium relationship. Implementing the Johansen test requires pre-specifying the VAR lag order. Following the suggestion by Hall (1989) and Johansen (1992), we specify the lag order such that the error terms are serially uncorrelated and the Akaike Information Criterion (AIC) is minimized.

Table 3
Unit root tests.

	Pre-crisis		Crisis	
	ADF	PP	ADF	PP
REIT index	−1.651	−1.756	0.391	0.490
	(0.771)	(0.725)	(0.999)	(0.999)
Δ REIT index	−20.814	−20.786	−19.536	−19.472
	(0.000)	(0.000)	(0.000)	(0.000)
Stock index	−1.571	−1.613	−0.230	−0.324
	(0.803)	(0.787)	(0.992)	(0.990)
ΔStock index	−22.483	−22.488	−20.645	−20.616
	(0.000)	(0.000)	(0.000)	(0.000)
Interest rate	−1.770	−1.932	−0.674	−0.784
	(0.718)	(0.636)	(0.974)	(0.965)
ΔInterest rate	−33.892	−38.243	−35.240	−37.090
	(0.000)	(0.000)	(0.000)	(0.000)

Notes: Δ indicates first order difference. ADF and PP denote the Augmented Dickey-Fuller and Phillips–Perron tests for stationarity, respectively. Both ADF and PP include a constant and a linear time trend. In ADF lag length is chosen by AIC. p-values are given in parantheses.

We proceed next with the specific threshold cointegration approach proposed by Enders and Granger (1998) and Enders and Siklos (2001). Unlike the Johansen (1988) method, which assumes linear behavior in the long and short run, this method allows for threshold adjustment in the short run while maintaining linearity in the long run. This two-step approach entails using ordinary least squares (OLS) to estimate the long-run equilibrium relationship as

$$p_{r,t} = \beta_0 + \beta_1 p_{s,t} + \beta_2 i_t + \mu_t \tag{1}$$

where $p_{r,t}$, $p_{s,t}$ and i_t are the natural logs of REIT index, stock market index and interest rate, respectively.[10] μ_t is the error term which may be serially correlated. To introduce asymmetric adjustment, Enders and Granger (1998) and Enders and Siklos (2001) let the deviation from the long-run equilibrium behave as a threshold autoregressive (TAR) process.

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1-I_t)\rho_2\mu_{t-1} + \sum_{w=1}^{p}\eta_w\Delta\mu_{t-w} + \varepsilon_t \tag{2}$$

where $\Delta\mu_{t-w}$ terms are added to account for serial correlation and I_t is the Heaviside indicator function[11] such that

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq \tau \\ 0 & \text{if } \mu_{t-1} < \tau \end{cases} \tag{3}$$

The Heaviside indicator depends on the level of μ_{t-1} and the threshold value, τ. An alternative specification, suggested by Enders and Granger (1998) and Caner and Hansen (1998), defines the Heaviside indicator to depend on the previous period's change in μ_{t-1}.

$$I_t = \begin{cases} 1 & \text{if } \Delta\mu_{t-1} \geq \tau \\ 0 & \text{if } \Delta\mu_{t-1} < \tau \end{cases} \tag{4}$$

This alternative specification is called momentum-threshold autoregressive (M-TAR) model. It is especially relevant when the adjustment is such that the series exhibits more momentum in one direction than the other. Whether TAR or M-TAR model is going to be used is based on a model selection criterion, such as AIC.

This framework presumes the value of the threshold is known; however in practice one has to estimate the value of the threshold. For this purpose, we use the methodology proposed by Chan (1993). This method eliminates the largest and smallest 15 percent of the estimated residual series, μ_t and considers each of the remaining 70 percent of the values as potential threshold. Using each value of potential threshold, equation (2) is estimated. The value yielding the lowest residual sum of squares is a consistent estimate of the threshold.

Once equation (2) is estimated based on the consistent estimate of the threshold, we compute the F-test of the null hypothesis of no cointegration (H_0:$\rho_1 = \rho_2 = 0$) versus the alternative of cointegration with TAR or M-TAR threshold adjustment. The critical values of this non-standard F-test are tabulated in Wane et al. (2004). If this null hypothesis is rejected, it is possible to test for the null hypothesis of symmetric adjustment (H_0:$\rho_1 = \rho_2$) by using standard F-test.

According to the Granger representation theorem, the existence of cointegration justifies estimating an error correction model. Given the existence of a cointegrating relationship in the form of (1), we estimate the following vector error-correction model:

$$\Delta p_{r,t} = \alpha_r + \delta_{r1}I_t\mu_{t-1} + \delta_{r2}(1-I_t)\mu_{t-1} + \sum_{q=1}^{k}\gamma_{rq}\Delta p_{s,t-q} + \sum_{q=1}^{k}\theta_{rq}\Delta p_{r,t-q} + \sum_{q=1}^{k}\eta_{rq}\Delta i_{t-q} + \upsilon_{r,t} \tag{5}$$

$$\Delta p_{s,t} = \alpha_s + \delta_{s1}I_t\mu_{t-1} + \delta_{s2}(1-I_t)\mu_{t-1} + \sum_{q=1}^{k}\gamma_{sq}\Delta p_{s,t-q} + \sum_{q=1}^{k}\theta_{sq}\Delta p_{r,t-q} + \sum_{q=1}^{k}\eta_{sq}\Delta i_{t-q} + \upsilon_{s,t} \tag{6}$$

$$\Delta i_t = \alpha_i + \delta_{i1}I_t\mu_{t-1} + \delta_{i2}(1-I_t)\mu_{t-1} + \sum_{q=1}^{k}\gamma_{iq}\Delta p_{s,t-q} + \sum_{q=1}^{k}\theta_{iq}\Delta p_{r,t-q} + \sum_{q=1}^{k}\eta_{iq}\Delta i_{t-q} + \upsilon_{i,t} \tag{7}$$

where $I_t\mu_{t-1}$ and $(1-I_t)\mu_{t-1}$ are the error correction terms, $\upsilon_{r,t}$, $\upsilon_{s,t}$ and $\upsilon_{i,t}$ are white noise errors and all variables are as defined above. The number of lags, k, is determined by both Ljung-Box statistic and AIC.

This threshold model allows the adjustment to depend on the deviation from the long-term equilibrium ($\mu_{t-1} \geq \tau$ vs $\mu_{t-1} < \tau$) for the TAR and on the change in the deviation from the long-term equilibrium ($\Delta\mu_{t-1} \geq \tau$ vs $\Delta\mu_{t-1} < \tau$) for the M-TAR process. M-TAR specification is especially relevant when the adjustment is such that the series exhibits more momentum in one direction than the other.

The coefficients in the vector error-correction model are used to tests for Granger causality. The hypothesis of the absence of Granger causality is tested separately for $\Delta\mu_{t-1} \geq \tau$ and $\Delta\mu_{t-1} < \tau$. Since the error correction terms include lagged levels of explanatory variables, we test the joint significance of the lags of differenced explanatory variable as well as the error correction term. For example, regarding causality from stock prices to real estate prices, we test:

$$\mathrm{H_0}: \gamma_{r1} = .. = \gamma_{rk} = \delta_{r1} = 0 \;\text{ for }\; \mu_{t-1} \geq \tau \;\text{ for TAR }\; (\Delta\mu_{t-1} \geq \tau \text{ for M}-\text{TAR}) \quad \text{and}$$

$$\mathrm{H_0}: \gamma_{r1} = .. = \gamma_{rk} = \delta_{r2} = 0 \;\text{ for }\; \mu_{t-1} < \tau \text{ for TAR } (\Delta\mu_{t-1} < \tau \text{ for M}-\text{TAR})$$

We examine the remaining cases (from stock to real estate market, from interest rate to real estate market etc.) in a similar way.

4. Empirical results

Once the order of integration of each variable is determined, we proceed to test for cointegration. The Johansen trace and maximum eigenvalue tests are applied to REIT index, stock market index and interest rate series separately during the pre-crisis and crisis periods. Table 4 presents the results. For the pre-crisis period both the trace and maximum eigenvalue statistics show the absence of

[10] Following Lean and Smyth (2012), we use an empirical specification that includes the natural logs of REIT index, stock market index, and interest rate.

[11] The Heaviside indicator function, or the unit step function, usually denoted by *H* is a discontinuous function whose value is zero for arguments that fall below a given threshold level and one otherwise. The function was named after Oliver Heaviside, who originally developed it in operational calculus.

cointegration. For the crisis period both statistics indicate the existence of one cointegration vector at the 10 percent significance level.

The results of both TAR and MTAR cointegration test are presented separately for the pre-crisis and crisis periods in Table 5. The table reports values of the adjustment coefficients ρ_1 and ρ_2, the F_{coint} statistics for the null hypothesis of no cointegration (i.e. $\rho_1 = \rho_2 = 0$) versus the alternative of cointegration with asymmetric adjustment as well as the $F_{symmetry}$ statistic for the null hypothesis of symmetric adjustment ($\rho_1 = \rho_2$). It also reports the estimated value of threshold, the lag length that is selected such that the Akaike Information Criterion (AIC) is minimized and Ljung–Box Q(k) statistic for the hypothesis that the first k of the residual autocorrelations are jointly equal to zero. The lower part of the table presents the estimation results for the long-run relationship given by Eq. (1).

As suggested by the F_{coint} statistics, cointegration is found neither during the pre-crisis period nor during the crisis period for the TAR model. This finding is consistent with the finding of the Johansen test, which could not detect cointegration during the pre-crises period and documents cointegration only marginally at 10 percent significance level for the crisis period. For the M-TAR model, however, the F_{coint} statistics show that cointegration exists during both periods. As can be seen from the AIC figures, the MTAR model shows a better fit to the data. Therefore, only the MTAR specification will be used in the remainder of the paper.

The necessary conditions for convergence are $\rho_1 < 0$, $\rho_2 < 0$ and $(1+\rho_1)(1+\rho_2) < 1$ (Petruccelli and Woolford, 1984). Considering the point estimates, one observes that for both periods ρ_1 but not ρ_2 is statistically significant and negative. It is clear that they satisfy the convergence conditions. The point estimates of ρ_1 and ρ_2 for both the pre-crisis and crisis periods suggest substantially faster convergence for the case$\Delta\mu_{t-1} \geq \tau$ than the case $\Delta\mu_{t-1} < \tau$. Moreover, the null hypothesis of symmetric adjustment is rejected at the 1% significance level in both periods.

After testing for threshold cointegration and for the equality of adjustment parameters, we proceed by formulating a threshold error correction model. We first focus on the coefficients of the error correction terms. Table 6 presents the estimated values and corresponding t-values of the two error correction terms, which correspond to above-threshold and below-threshold cases. Note that the signs of these coefficients should be interpreted in conjunction with the signs of the coefficients in equation (1), which gives the long run equilibrium. As is reported in the lower part of Table 5, the coefficient of stock price index is positive in both periods. On the other hand, the coefficient of interest rate is negative in the pre-crisis. It becomes positive in the crisis period. Based on these signs, to eliminate deviations from long run equilibrium the coefficients of error correction terms in equation (5) (δ_{r1} and δ_{r2}) should be negative and those in equation (6) (δ_{s1} and δ_{s2}) should be positive in both periods. On the other hand, the coefficients of error correction terms in equation (7) (δ_{i1} and δ_{i2}) should be negative in the pre-crisis and positive in the crisis period.

As can be seen from Table 6, adjustment to eliminate deviations from long-run equilibrium occurs only when $\Delta\mu_{t-1} \geq \tau$ in both periods. For the pre-crisis period, both real estate prices and interest rates adjust to eliminate deviations.[12] For the crisis period, on the other hand, it is mainly stock prices that adjust to eliminate deviations. F tests for the equality of the error correction terms

Table 4
Johansen cointegration test.

	Trace			Maximum Eigenvalue			Lag
	r = 0	r ≤ 1	r ≤ 2	r = 0	r ≤ 1	r ≤ 2	
Pre-crisis	22.965	4.615	0.130	18.350	4.485	0.130	1
	(0.248)	(0.848)	(0.718)	(0.117)	(0.804)	(0.718)	
Crisis	40.846	17.720	4.977	23.126	12.744	4.977	7
	(0.080)	(0.363)	(0.600)	(0.101)	(0.349)	(0.600)	

Notes: p-values are provided in parentheses. Lag order is chosen such that the error terms are serially uncorrelated and the Akaike Information Criterion (AIC) is minimized.

Table 5
Threshold cointegration tests.

	Pre-crisis		Crisis	
	TAR	MTAR	TAR	MTAR
ρ_1	−0.016	−0.210[a]	−0.045 [b]	−0.130[a]
	(0.817)	(7.919)	(2.414)	(4.596)
ρ_2	−0.036[b]	0.014	−0.033[c]	−0.017
	(2.078)	(0.993)	(1.955)	(1.241)
F_{coint}	2.442	31.910[a]	4.779	11.287[a]
$F_{symmetry}$	0.552	56.329	0.238	13.014
	(0.458)	(0.000)	(0.626)	(0.000)
τ	−0.070	0.513	0.035	0.017
Lag	4.000	2	1.000	1
Q(5)	0.632	4.259	1.664	1.591
AIC	−5.261	−5.366	−4.956	−4.981
	Cointegration relationship			
β_0	5.981		−4.355	
	(23.816)		(32.836)	
β_1	0.543		1.351	
	(24.032)		(108.450)	
β_2	−0.462		0.027	
	(24.234)		(2.184)	

Notes: ρ_1 and ρ_2 are the adjustment parameters corresponding to the above threshold and below threshold cases, respectively. F_{coint} tests the null hypothesis of no cointegration (H_0:$\rho_1 = \rho_2 = 0$) and $F_{symmetry}$ tests the null hypothesis of symmetric adjustment (H_0:$\rho_1 = \rho_2$). τ denotes the estimated threshold. Lag denotes optimal lag length chosen by Akaike Information Criterion (AIC). Q(5) is the Ljung–Box statistic for the hypothesis that the first 5 of the residual autocorrelations are jointly equal to zero. For ρ_1 and ρ_2 t-statistics are given in parantheses. F_{coint} follows a non-standard distribution.The critical values of this test are tabulated in Wane et al. (2004). $F_{symmetry}$ follows standard F distribution. For $F_{symmetry}$ p-values are given in parantheses. For 10% significance level, critical values for Q(5) is 9.236. For the coefficients of the cointegration relationship t-statistics are given in parantheses. Note that the coefficient estimates are superconsistent and they do not have asymptotic t-distribution. [a],[b], and [c] indicate significance at 1%, 5% and 10% level, respectively.

Table 6
Error correction terms for the MTAR model.

Dependent	Pre-crisis			Crisis		
	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$F_{equality}$	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$F_{equality}$
ΔREIT Index	−0.131[a]	0.030 [b]	12.182	0.024	0.007	0.522
	(4.508)	(2.025)	(0.000)	(0.862)	(0.548)	(0.594)
ΔStock Index	−0.061[b]	0.017	2.836	0.115[a]	0.018	6.859
	(2.102)	(1.125)	(0.060)	(3.525)	(1.134)	(0.001)
ΔInterest Rate	−0.256[a]	−0.020	17.326	−0.024	0.003	0.484
	(5.816)	(0.888)	(0.000)	(0.941)	(0.287)	(0.617)
AIC		−15.852			−15.884	
Lag		1			1	

Notes: The table reports the estimated coefficients of the two error correction terms (corresponding to above-threshold and below-threshold cases) separately for the pre-crisis and crisis periods. $F_{equality}$ tests the null hypothesis that these coefficients are equal to each other. For the estimated coefficients of error correction terms t-statistics are given in parantheses. For $F_{equality}$ test p-values are given in parantheses. AIC denotes Akaike Information Criterion and Lag shows the optimal lag length chosen by AIC. [a],[b], and [c] indicate significance at 1%, 5% and 10% level, respectively.

[12] For the pre-crisis period, real estate prices (stock prices) adjust in the wrong direction when $\Delta\mu_{t-1}<\tau$ ($\Delta\mu_{t-1}\geq\tau$).

Table 7
Granger causality results: MTAR model pre-crisis period.

Panel A	ΔStock index		ΔInterest rate	
	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$
	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$
ΔREIT Index	10.171 (0.000)	2.095 (0.124)	10.413 (0.000)	2.072 (0.127)
Panel B	ΔREIT index		ΔInterest rate	
	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$
	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$
ΔStock Index	2.932 (0.054)	1.875 (0.155)	2.370 (0.095)	0.643 (0.526)
Panel C	ΔREIT Index		ΔStock Index	
	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$
	$\Delta\mu_{t-1}\geq\tau$	$\Delta\mu_{t-1}<\tau$	$\Delta\mu_{t-1}\geq\tau$	$\Delta\mu_{t-1}<\tau$
ΔInterest Rate	18.883 (0.000)	1.464 (0.232)	17.313 (0.000)	0.917 (0.401)

Notes: $F_{\text{No causality}}$ test the joint significance of the lags of differenced explanatory variable as well as the error correction term separately for the above-threshold and below-threshold cases. For $F_{\text{No causality}}$ test p-values are given in parantheses.

Table 8
Granger causality results: MTAR model crisis period.

Panel A	ΔStock index		ΔInterest rate	
	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$
	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$
ΔREIT index	0.544 (0.581)	0.319 (0.727)	0.505 (0.604)	0.260 (0.771)
Panel B	ΔREIT index		ΔInterest rate	
	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$
	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$
ΔStock Index	7.272 (0.001)	1.599 (0.203)	7.063 (0.001)	1.274 (0.281)
Panel C	ΔREIT Index		ΔStock Index	
	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$	$F_{\text{No causality}}$
	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$	$\Delta\mu_{t-1} \geq \tau$	$\Delta\mu_{t-1} < \tau$
ΔInterest Rate	0.837 (0.434)	0.383 (0.682)	0.446 (0.641)	0.041 (0.960)

Notes: $F_{\text{No causality}}$ test the joint significance of the lags of differenced explanatory variable as well as the error correction term separately for the above-threshold and below-threshold cases. For $F_{\text{No causality}}$ test p-values are given in parantheses.

indicate that while stock price index displays asymmetry in both periods, REIT index and interest rates display asymmetry only during the pre-crisis period.

To further investigate the dynamic behavior of stock and real estate prices, we examine next the nature of the Granger causality between REITs, stock prices and interest rates using the estimation results of threshold error correction model. The results of the causality tests for the pre-crisis period are presented in Table 7. We focus on panels A and B, in which the dependent variable is REIT index and stock index, respectively. We draw conclusion based on two F statistics, which test for the absence of causality separately for $\Delta\mu_{t-1} \geq \tau$ and $\Delta\mu_{t-1} < \tau$. The evidence provides support for the existence of both credit-price and wealth effects. It is notable that only the F statistic associated with the above threshold regime ($\Delta\mu_{t-1} \geq \tau$) is significant when either real estate prices or stock prices is the dependent variable. Moreover, interest rate does Granger cause both real estate and stock prices.

The results of the causality tests for the crisis period are presented in Panels A and B of Table 8. When real estate price is the dependent variable, the insignificance of the two F tests suggests the absence of Granger causality from stock prices to real estate prices. On the other hand, when stock price is the dependent variable, the F statistics associated with the above threshold regime ($\Delta\mu_{t-1} \geq \tau$) is significant, indicating the existence of Granger causality from real estate prices to stock prices. Overall, these results show that credit-price effect but not wealth effect exists during the crisis period. Moreover, interest rate does Granger cause stock but not real estate prices.

To complement the picture panel C of Tables 7 and 8 give the results when interest rate is the dependent variable. The findings indicate that during the pre-crisis period both real estate and stock prices Granger cause interest rates. In contrast, neither variable Granger causes interest rates during the crisis period.

Overall, when we compare pre-crisis and crisis periods, it is noteworthy that while both credit-price and wealth effects exist during the pre-crisis period, only a credit-price effect is observed during the crisis period.[13] The existence of both effects during the pre-crisis period is consistent with the finding in Lean and Smyth (2012) for Malaysia, Liu and Su (2010) for China and Su (2011) for France, Spain and Switzerland.

The major finding in this study is the disappearance of the wealth effect in the crisis period. Based on the earlier discussion of the economic conjuncture, it seems possible that weak domestic and external demand during the crisis period has lead to a structural change in the behavior of households. This may have eliminated the previously existing impact of wealth on households' demand for real estate.

5. Conclusions

This study investigates the extent to which the dynamic relationship between real estate prices and stock prices in Turkey has changed as a result of the global financial crisis. Our analyses include interest rate, which is likely to have an impact on investors' ability to finance investments in real estate and stock markets (Chen, 2001), as a control variable. The findings are as follows. First, our evidence gives empirical support to the argument in Gonzalo and Pitarakis (2006) that omitting the presence of nonlinear components, like threshold effects in long-run equilibrium, may lead to wrong inferences regarding the existence of cointegration. In the analyses, we employ both the Johansen test, which implicitly assumes linear error correction mechanism and the momentum threshold cointegration test that allows for asymmetric error correction. Unlike the Johansen test, which finds cointegration only during the crisis period, the momentum threshold test identifies cointegration in both pre-crisis and crisis periods.

Second, there is evidence that the nature of the long-run relation between these three variables has changed after the global financial crisis. When we compare pre-crisis and crisis periods, it is noteworthy that while both credit-price and wealth effects exist during the pre-crisis period, only a credit-price effect is observed during the crisis period. One possible explanation for the disappearance of the wealth effect is the change in economic conjuncture after the outbreak of the crisis.

Our analysis also detects other differences between the pre-crisis and crisis periods. We find that both real estate prices and

[13] To check the robustness of our results to the choice of the starting date of the crisis, the analyses are repeated by choosing October 1, 2007 and November 1, 2007 as the starting dates. The results (not reported) are qualitatively the same and are available from the author upon request.

interest rates adjust to eliminate deviations during the pre-crisis period. This role is played by stock prices during the crisis period. Finally, while interest rates Granger cause both real estate and stock prices in pre-crisis period, they Granger cause only stock prices in the crisis period.

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