A Comparison of Currency Crisis Dating Methods: Turkey 1990-2014

ALI ARI[1], and RAIF CERGIBOZAN[2]

[1] Department of Economics, Kirklareli University, Turkey, E-mail: ali.ari@klu.edu.tr
[2] Department of Economics, Kirklareli University, Turkey, E-mail: raif.cergibozan@klu.edu.tr

ARTICLE INFO

ABSTRACT

This paper aims to assess the robustness of currency crisis dating methods. Hence, we reproduce a broad set of the ten most representative indicators from the literature, and develop two new crisis indicators derived from Cointegration and ARDL Bounds tests for the Turkish economy which underwent several currency crises over the last 30 years. Contrary to early studies (Edison, 2003; Perez, 2005; Lestano and Jacobs, 2007), we indicate that different crisis indicators produce similar results, at the same threshold level, in identifying the Turkish currency crises of post-liberalization period.

INTRODUCTION

The last three decades has seen the proliferation of currency crises in several developed and emerging economies. This naturally led to the development of a large number of theoretical and empirical studies in order to understand crisis mechanisms. The empirical papers that aim at identifying crisis determinants in order to predict future crises are frequently called "early warning systems" (EWS), as they are likely to inform policymakers and investors about the occurrence of a crisis in the near future.

Identifying crisis episodes is crucial for the empirical papers. However, different crisis definitions are developed in the literature that may lead to significant discrepancies in crisis dates.

Moreover, using different crisis indicators may also change the significance of explicative variables in model estimations. Hence, this paper aims to assess the robustness of crisis indicators in the signalling of crises. For doing so, we reproduce ten different currency crisis indicators, frequently used in the literature, and develop two new crisis indicators, derived from Johansen Cointegration method and ARDL test. This new approach enables us to solve the weighting problem, which is one of the main drawbacks of the existing crisis indexes, since the weights of the index components are endogenously assigned by our model estimations. On the other hand, this approach which shows whether there is a long-run relationship between the index components allows us to evaluate the robustness of the existing approaches in crisis dating literature.

We use data on Turkey for the post-liberalization era (January 1990-June 2014). Turkey which underwent several currency crises (1991, 1994, 2001, 2006, 2008, and 2013) over the last thirty years is an interesting case study.[1] This single-country approach also allows us to better assess the performance of different crisis indexes in crisis dating. The outcomes of this study may then be extrapolated to other country samples.

There actually exist two distinct periods in the Turkish economy in its post-liberalization era: before and after the 2001 financial crisis. During the pre-2001 crisis period, Turkey presented important macroeconomic imbalances (unstable economic growth, excessive budget deficits, high and chronic inflation, significant dollarization, sharp rises in short-term external debt), significant banking sector weaknesses (high non-performing loans, important currency mismatches, low own equity and reserves), and it also underwent some external shocks mostly related to the 1998 Russian crisis. This is why one can affirm that the crises occurred in that period mostly result from domestic problems.[2]

In the post-2001 crisis period, Turkey reached to restore a relative economic and financial stability with a radical structural economic program backed by the IMF stand-by credits. However, the country suffered a severe crisis in 2008-09, and high speculative pressures in 2006 and 2013, mostly related to external economic and financial problems. The economic growth strategy, which is heavily dependent on export earnings and short-term capital inflows, seems the main reason behind the external vulnerability of the country.

This paper has some common points with Lestano and Jacobs (2007)' study, but goes beyond it as it assesses the performance of much more crisis indicators in a more systematic way. Moreover, developing less-arbitrary crisis indicators through Cointegration and ARDL methods, firstly introduced by this paper, is an important contribution to the existing crisis literature.

The paper is organized as follows. Section 2 discusses in detail the currency crises dating schemes. Section 3 presents the Turkish data, currency crisis indicators and the results of different crisis indicators. Section 4 concludes.

2. DATING CRISES: A LITERATURE SURVEY

Identifying crisis dates is crucial to all empirical studies. Some papers describe currency crises as large depreciation or devaluation episodes (Frankel and Rose, 1996; Milesi-Ferretti and Razin, 2000; Glick and Moreno, 1999; Esquivel and Larrain, 2000; Osband and Van Rijckeghem, 2000; Kumar et al., 2003, Gourinchas and Obstfeld, 2011).[3] The question here is how big a change in the exchange rate is qualified as a currency crisis. Thus, all papers determine

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[1] See Nas (2008) and Ari and Cergibozan (2014) for a detailed analysis on the recent financial crises occurred in the Turkish economy.

[2] Note also that these domestic problems were worsened by highly volatile capital in- and outflow following the financial liberalization through 1980s in Turkey.

[3] As many countries use the US dollar to define their exchange rate; nearly all papers employ nominal or real bilateral dollar exchange rates for constructing crisis indicators.

an arbitrary threshold; if the indicator value exceeds this specified threshold, any month, quarter or year is classified as a crisis episode.

For Frankel and Rose (1996), an annual depreciation of at least 25% is considered as currency crisis. For the countries with high inflation rates and correspondingly high rates of depreciation, the authors put another criterion to not consider each of these depreciations as an independent crisis episode: the change in the exchange rate should also exceed the previous year's change in the exchange rate by a margin of at least 10%.

For Glick and Moreno (1999), a crisis occurs when the monthly percentage change in the exchange rate exceeds the index mean (μ) plus two standard deviations ($2 \times \sigma$). For Esquivel and Larrain (2000), a currency crisis occurs if the accumulated three-month real exchange rate is greater than 15% or the one-month real exchange rate is higher than 4% and also greater than the index mean (μ) plus 2.54 standard deviations ($2.54 \times \sigma$).

Others define currency crises as cases where a currency comes under severe speculative pressure. This second definition takes into account both the situations where speculative attacks lead to depreciation and where the authorities successfully defend the currency by selling foreign reserves and/or rising domestic interest rates. This speculative pressure is measured by an index (*ISP*, index of speculative pressure or *EMPI*, exchange market pressure index) that is as a weighted average of changes in nominal or real exchange rate[4] (*NER* or *RER*) and changes in foreign reserves (*RES*) (Kaminsky et al., 1998; Cartapanis et al., 1998; Kaminsky and Reinhart, 1999; Berg and Pattillo, 1999a, 1999b; Bussiere and Mulder, 2000; Vlaar, 2000; Kamin et al., 2001; Edison, 2003; Komulainen and Lukkarila, 2003; Caramazza et al., 2004; Alvarez-Plata and Schrooten, 2004; Budsayaplakorn et al., 2010; Licchetta, 2011; Aizenman and Pasricha, 2012; Ari and Cergibozan, 2016). Some others also include changes in nominal or real interest rates (*NIR* or *RIR*) into their crisis indexes (Eichengreen et al., 1996; Hawkins and Klau, 2000; Bussiere and Fratzscher, 2006).

The weights of the components of the crisis index are chosen to equalize their volatility and thus avoid the possibility of one of the components dominating the index (Aziz et al., 2000). Note that the weights are frequently the inverse of the standard deviation (σ) of the corresponding component. However, Bussiere and Fratzscher (2006) and Bussiere (2013) prefer to use the inverse of the variance (σ^2) of the corresponding component. On the other hand, some authors also utilize arbitrary weights: for instance in Corsetti et al. (2001) the weights assigned to exchange rate and reserves are, respectively, 0.75 and 0.25. Herrera and Garcia (1999) and Krkoska (2001) use standardized weights for index components.

Once components of the crisis index and their weights are determined, an arbitrary value of threshold is specified above which the index signals a crisis. The crisis index then becomes a binary variable which takes the value of 1 if a crisis occurs and 0 otherwise. The threshold values are generally set to a multiple of the standard deviation of the index (σ_{ISP}) plus the index mean (μ_{ISP}). The values of thresholds used in the literature have been ranged from 1 to 3 standard deviations.

Many authors using the first or second approach generally tend to incorporate exclusion windows in order not to illustrate the same crisis more than once. For those authors, currency crisis episodes are identified rather than currency crises, since a currency crisis can reveal itself

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[4] Since a large nominal depreciation may simply reflect the repercussion of high inflationary episodes but not the occurrence of speculative attacks, many empirical studies correct this bias by either defining the behaviour of the exchange rate in real terms (Cartapanis et al., 1998; Bussiere and Fratzscher, 2006; Ari, 2012) or by dividing the sample into low and high inflationary periods with corresponding different estimations (Kaminsky et al., 1998; Herrera and Garcia, 1999; Glick and Moreno, 1999).

through many crisis events (Lestano and Jacobs, 2007). This exclusion window may be in the range of 3 months to 3 years.

Both approaches have strengths, but also weaknesses. The devaluation approach may be criticized for its limited crisis definition given that every speculative attack leads to some economic cost even if the attack is successfully defended by the government. Thus, this methodology mostly captures currency crashes, but not necessarily currency crises as pointed out by Frankel and Rose (1996) and Pozo and Amuedo-Dorantes (2003).

However, this approach might be justified, because speculative attacks lead in general to depreciation of the currency, particularly in emerging market economies. Hence, including foreign reserves and/or interest rates may not have any effect on the crisis indicator. Moreover, for some authors, including foreign reserves and/or interest rates may lead to a misleading interpretation of crisis. Frankel and Rose (1996) consider that reserve movements are notoriously noisy measures of exchange market intervention for almost all countries. Glick and Moreno (1999) explain this issue by the fact that government interventions in currency markets may be used for debt payments, not only to defend the currency.

Besides, Flood and Marion (1998) and Frankel and Saravelos (2010) indicate that reserves stock may show large jumps at the peak of the crisis if central banks draw credit under IMF programs or from other foreign institutions. This may reduce the impact of changes in exchange rate over the index. As a consequence, the index may fail to detect certain crisis incidences like in Philippines in September 1997, as shown by Zhang (2001). Furthermore, Frankel and Saravelos (2010) also point out that movements in exchange rates cause severe valuation distortions in reserves stock. Hence, the drop in foreign reserves may be considered as a biased measure of crisis incidence.

On the other hand, the inclusion of interest rates into the crisis index is criticized for two reasons. First, some authors like Kaminsky and Reinhart (1999) argue that the data on market-determined interest rates in developing countries is incomplete and inadequate. Second, the impact of rising domestic interest rates on speculative attacks is controversial (Arias, 2003). As well-documented in second generation crisis models, a rise in interest rates leads to increasing concerns of economic agents on the sustainability of the exchange rate, particularly in a highly indebted economy, instead of calming down financial markets (Ari, 2012).

However, selling reserves and rising interest rates are common practical tools used by monetary authorities to face speculative attacks. This defence policy may or may not stabilize the exchange rate. Therefore, it is justified to consider currency crises as high pressures on the currency. Besides, the inclusion of reserves and/or interest rates as a measure of crisis incidence allows one to observe an increase in market pressure that may not be captured through changes in exchange rate. This is particularly meaningful for countries that adopted fixed exchange rate regime, where capital outflows and crisis are manifest through larger drops in reserves and/or increasing interest rates rather than exchange rate weaknesses (Frankel and Saravelos, 2010).

One can notice that this empirical definition is rather independent from a structural model of exchange rates even though Eichengreen et al. (1994) were inspired by Girton and Roper (1977)'s monetary model.[5] The idea here is that the index should ideally stem from a structural model of exchange rate determination (Arias, 2003). However, since the empirical performance of such structural models is relatively poor (Meese and Rogoff, 1983), the focus is rather on the origin of successful or unsuccessful speculative attacks to justify the inclusion of stabilizing measures for the exchange rate in the empirical definition of the index.

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[5] This ISP or EMPI is derived by Eichengreen et al. (1994, 1995, 1996) from the monetary model of Girton and Roper (1977). This model shows that a change in both international reserves and prices of foreign exchange can affect the foreign exchange market or the pressure on the external position of a country.

The speculative pressure approach is mostly criticized because of arbitrary weighting procedure of the index components. As underlined by Eichengreen et al. (1994), the simplest solution is to construct a crisis index with equal or normalized weights. But, the volatility of the index components is different. Thus, this 'simple' solution is not 'optimal' even if it is used by some authors like Herrera and Garcia (1999) and Krkoska (2001).

With no theoretical guidance on choosing the 'optimal' weights, the standard procedure is to divide the index components by their standard deviation in order to equalize their volatility. However, this procedure cannot solve the volatility problem over time, as the relative variance of changes in exchange rate, reserves, and interest rates would be heavily influenced by the adopted exchange rate regime and the government's reaction function (Nitithanprapas and Willett, 2002). In the fixed (floating) exchange rate regime, changes in international reserves (exchange rate) are more common than changes in exchange rate (international reserves). Therefore, using this weighting procedure over a period that spans two different exchange rate regimes would result in information loss, hence bias the index (Chui, 2002).

To avoid this volatility problem, Aziz et al. (2000), Nitithanprapas and Willett (2002) and Caramazza et al. (2004) propose to deseasonalize or detrend the data of the index components. Moreover, Eichengreen et al. (1995, 1996) and Nitithanprapas and Willett (2002) also suggest testing different weights for sensitivity issues. But they also affirm that different weights do not change significantly dating of currency crisis episodes.

The most important problem related to crisis identification by the construction of crisis indexes (both approaches) is the definition of arbitrary values of threshold above (below) which the observation is classified as a 'crisis' ('non-crisis') with the associated unit (zero) value. Note that higher the threshold level is the less the number of detected crisis is, and vice versa. Therefore, this arbitrary threshold method results in different number and effective dates of currency crises as empirically shown by Kamin et al. (2001), Edison (2003), and Lestano and Jacobs (2007).

On the other hand, Zhang (2001) and Edison (2003) state that values of threshold for ISP indexes are sample-dependent. Hence, when a severe crisis occurs which leads to large movements in exchange rates, the sample mean and the sample standard deviation can change substantially, causing changes in crisis dates. In other words, the sample-dependent nature of threshold definition implies that future data can affect the identification of past crisis (Abiad, 2003).

Note also that crisis dates may change from one study to another according to the index components (whether interest rates are included into the index), and to the nominal or real character of the index components (Ari, 2012). In addition, Chui (2002) refers to another problem concerning data: utilizing end-period figures or period-average figures in the index calculation where the sample standard deviation of period-average data is generally smaller and resulting in a lower crisis threshold might slightly change the identification of crisis dates. Moreover, Reagle and Salvatore (2005) point out another data problem related to data revisions routinely introduced by the international organizations, i.e. the IMF or the World Bank. For the authors, this can lead to significant revisions in empirical results as they showed in predicting the 1997 Asian crisis.

The last problem with the currency crisis dating schemes is the exclusion window which omits any crises identified by the threshold method if they follow a previous crisis within a certain length of time (Abiad, 2003). But this *ad hoc* adjustment introduces artificial serial correlation in the dependent variable and eliminates any information on crisis duration.

3. DATA AND CRISIS INDEXES
3.1 Data

The discussion on the dating methodology indicates no system could perfectly track crisis episodes. Thus, combining the results from the quantitative definition with country-specific knowledge about exchange market developments would improve the analysis (Berg et al., 2005). Using different crisis definitions and different values of threshold in order to assess the robustness of the crisis dating schemes is another solution, like we do in this paper.

The study focuses on the Turkish economy and covers Turkey's entire post-liberalization era from January 1990 to June 2014. The monthly data is gathered from the IMF's International Financial Statistics. Exchange rates used to calculate the indexes are defined in terms of US dollars and market rates, international reserves are measured as total reserves minus gold and in US dollars, and domestic interest rate is 3-month deposit rate.

3.2 Crisis Indicators

This paper constructs twelve currency crisis indicators for the Turkish economy; the first ten are the replication of the most commonly used indicators in the literature, while the last two are derived from Cointegration and ARDL Bounds tests. This approach which shows whether there is a long-run relationship between the index components, i.e. exchange rate, international reserves, and domestic interest rates, allows us to evaluate the robustness of the existing approaches in crisis dating literature. If there is a long-run relationship between the index components, this will rather confirm the ISP approach; if there is no long-run relationship between the index components, this will rather confirm the depreciation approach.

Besides, this method also allows us to solve the weighting problem, which is one of the main drawbacks of the existing crisis indexes. As underlined in the previous section, the index components are generally weighted by the inverse of their standard deviation or variance in order to equalize their volatility. However, equalizing volatility of the index components leads to loss of information, since the high volatility is generally observed just before or during crisis episodes. In the Cointegration and ARDL tests, the weighting procedure is realized through long-run estimations that avoids the loss of information.

On the other hand, we use four values of threshold (β) in terms of standard deviation (1, 1.5, 2, 3) above the index mean to illustrate the sensitivity. Note that contrary to original papers, we do not use exclusion window for any of our crisis indicators.

3.2.1 Eichengreen et al. (1996)

This index is constructed as a weighted average of changes in nominal exchange rate, changes in international reserves, and changes in nominal interest rates. All of these variables are measured relative to those prevailing in the US, the reference country, and weighted by the inverse of their respective standard deviation.

$$ISP1_t = \frac{1}{\sigma_{NER}} \Delta NER_{TR,t} - \frac{1}{\sigma_{RES}} \left(\Delta RES_{TR,t} - \Delta RES_{US,t} \right) + \frac{1}{\sigma_{NIR}} \Delta \left(NIR_{TR,t} - NIR_{US,t} \right) \tag{1}$$

where NER_{TR} is nominal exchange rate, RES_{TR} and RES_{US} are foreign reserves for Turkey and the US, respectively, NIR_{TR} and NIR_{US} are nominal interest rates for Turkey and the US, respectively, σ_{NER}, σ_{RES}, and σ_{NIR} are the standard deviations of the index components.

A crisis is identified when the index exceeds some upper bound:

$$Crisis_t = \begin{cases} 1 & \text{if } ISP1_t > \beta \times \sigma_{ISP1} + \mu_{ISP1} \\ 0 & \text{otherwise} \end{cases} \tag{2}$$

where β is the value of threshold, σ_{ISP1} and μ_{ISP1} represent the index standard deviation and the index mean, respectively.

3.2.2 Kaminsky et al. (1998)

This index is composed of changes in nominal exchange rate ΔNER and changes in foreign reserves ΔRES. Compared to the previous indicator, the authors drop the links to the reference country and interest rate differential, and multiply changes in reserves by the standard deviation of the relative change in the exchange rate. By doing this, they aim to equalize the volatility of reserves to the volatility of exchange rate.

$$ISP2_t = \Delta NER_t - \frac{\sigma_{NER}}{\sigma_{RES}} \Delta RES_t \tag{3}$$

where σ_{NER} and σ_{RES} are the standard deviations of changes in nominal exchange rate and changes in foreign reserves, respectively.

3.2.3 Corsetti et al. (2001)

This index is composed of changes in nominal exchange rate ΔNER and changes in international reserves ΔRES. The weights assigned to exchange rate and reserves are, respectively, 0.75 and 0.25.

$$ISP3_t = 0.75\Delta NER_t - 0.25\Delta RES_t \tag{4}$$

3.2.4 Krkoska (2001)

This unweighted speculative pressure index is composed of changes in nominal exchange rate ΔNER, changes in international reserves ΔRES, and changes in nominal interest rates $(NIR_t - NIR_{t-1})$. Putting forward the empirical proofs of Eichengreen et al. (1996) and Corsetti et al. (2001) that different weights do not have any significant impact on the crisis index, Krkoska (2001) opts for the 'simplicity' and uses standardized weights for the index components.

$$ISP4_t = \Delta NER_t - \Delta RES_t + \left(NIR_t - NIR_{t-1} \right) \tag{5}$$

3.2.5 Von Hagen and Ho (2003)

This index is a weighted average of changes in real exchange rate, changes in reserves, and changes in nominal interest rate. The weights are the inverse of the standard deviation of each component. The authors use nominal instead of real interest rates, because nominal rates are what the authorities can directly control in order to ward off speculative attacks. Moreover, they

use the real exchange rate instead of nominal exchange rate, since periods of nominal depreci-ations due to inflation differentials should not be considered as currency crises.

$$ISP5_t = \frac{1}{\sigma_{RER}}\Delta RER_t - \frac{1}{\sigma_{RES}}\Delta RES_t + \frac{1}{\sigma_{NIR}}\left(NIR_t - NIR_{t-1}\right) \qquad (6)$$

where RER denotes real exchange rate, and is calculated as follows: $RER = (NER \times P^*)/P$, with NER is nominal bilateral TL/US dollar exchange rate, P^* and P are consumer price in-dexes in the US and Turkey, respectively, RES denotes foreign reserves, NIR is nominal in-terest rate, σ_{RER}, σ_{RES} and σ_{NIR} are the standard deviations of the index components.

3.2.6 Bussiere and Fratzscher (2006)

This index is composed of real exchange rate changes, reserves changes, and real interest rate changes. The weights are the relative precision or variance (which is equal to the square of standard deviation) of each variable so as to give a larger weight to the variables with less vola-tility.

$$ISP6_t = \frac{1}{\sigma^2_{RER}}\Delta RER_t - \frac{1}{\sigma^2_{RES}}\Delta RES_t + \frac{1}{\sigma^2_{RIR}}\left(RIR_t - RIR_{t-1}\right) \qquad (7)$$

where RER denotes real exchange rate and calculated as above, RES denotes foreign re-serves, RIR is real interest rate which is obtained by the simple Fisher equation: $RIR = NIR - \pi$ with $NIR =$ nominal interest rate and $\pi =$ inflation rate. σ^2_{RER}, σ^2_{RES} and σ^2_{RIR} are the variance of the index components.

3.2.7 Candelon et al. (2012)

The index used in Yiu et al. (2009) and Candelon et al. (2012) is an augmented version of Kaminsky et al. (1998)'s index. It is constructed as a weighted average of changes in nominal exchange rate ΔNER, changes in reserves ΔRES, and changes in nominal interest rate $(NIR_t - NIR_{t-1})$. The weight of each component is normalized by the sample standard deviation of the nominal exchange rate σ_{NER}.

$$ISP7_t = \Delta NER_t - \frac{\sigma_{NER}}{\sigma_{RES}}\Delta RES_t + \frac{\sigma_{NER}}{\sigma_{NIR}}\left(NIR_t - NIR_{t-1}\right) \qquad (8)$$

3.2.8 Alvarez-Plata and Schrooten (2006)

This index is constructed as a weighted average of nominal exchange rate changes ΔNER, foreign reserves changes ΔRES, and changes in interest differential between Turkey and the US $\Delta(NIR_{TR} - NIR_{US})$. The weights are calculated as the inverse of the series' standard devia-tion.

$$ISP8_t = \frac{1}{\sigma_{NER}}\Delta NER_{TR,t} - \frac{1}{\sigma_{RES}}\Delta RES_{TR,t} + \frac{1}{\sigma_{NIR}}\Delta\left(NIR_{TR,t} - NIR_{US,t}\right) \qquad (9)$$

where σ_{NER}, σ_{RES} and σ_{NIR} are the standard deviation of the index components.

3.2.9 Bunda and Ca'Zorri (2010)

Following Sachs et al. (1996), Bunda and Ca'Zorri (2010) define an index as the weighted average of three-month changes in nominal exchange rate NER and three-month changes in reserves RES. The index components are in natural logarithmic form and the inverse of the variance of the index components (σ^2_{NER} and σ^2_{RES}) is chosen as weighting procedure.

$$ISP9_t = \ln\left(\frac{NER_t}{NER_{t-3}}\right) - \frac{\sigma^2_{NER}}{\sigma^2_{RES}} \ln\left(\frac{RES_t}{RES_{t-3}}\right) \tag{10}$$

3.2.10 Glick and Moreno (1999)

This crisis indicator which represents the depreciation approach takes into account only real exchange rate changes RER which is defined as in Von Hagen and Ho (2003) or in Bussiere and Fratzscher (2006).

$$ISP10 = \Delta RER_t \tag{11}$$

3.2.11 The crisis indicator derived from Johansen Cointegration test

The eleventh crisis indicator is derived from Johansen Cointegration test. For the Johansen procedure, we initially determine the rank of the long-run matrix Π, which permits us to find the number of linearly independent columns of Π. This gives us the number of cointegrating relationships (vectors) that exist among variables. There are two test statistics for the number of cointegrating vectors: the trace test and the maximum eigenvalue statistics. In the trace test, the null hypothesis is that the number of cointegrating vectors is less than or equal to r, where r is 0, 1, or 2. In other words, the trace statistic tests the null of $r = k(k = 1, 2, ..., n-1)$ against the alternative of unrestricted r. The maximum eigenvalue statistic, on the other hand, tests that there are r cointegrating vectors against the alternative that $r+1$ exists.[6]

Before estimating cointegration test, we need first to check integrating properties of variables. We use the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests to test the stationarity of variables. According to the results, all variables are stationary at their first difference. In other words, they are integrated of order one I(1). Since all variables in our model are integrated of the same order, we then proceed to test the following equation for cointegration by utilizing the Johansen procedure:

$$z_t = [RER \ RES \ NIR] \tag{12}$$

where RER denotes real exchange rate, RES denotes international reserves, and NIR denotes nominal interest rate.

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[6] See Johansen (1988), Johansen and Juselius (1990) and Johansen (1995) for further details on cointegration test.

Table 1. Johansen's test for number of cointegrating vectors

	Test Statistics			
	Trace		Maximum Eigenvalue	
	Null Statistic	5% C.V.	Null Statistic	5% C.V.
r =0	31.12395**	24.27596	20.98562**	9.474804
r ≤ 1	10.13833	12.32090	17.79730	2.976163
r ≤ 2	0.654446	4.129906	0.654446	4.129906

Test results are given in Table 1. According to both test statistics suggested by Johansen and Juselius (1990), there exists a single cointegrating (long-run) relationship among the variables of the model. The existence of a long-run relationship constitutes our eleventh crisis index, and also confirms the speculative pressure approach. We set then four values of threshold above the index mean to obtain crisis dates for the Turkish economy.

3.2.12 The crisis indicator derived from ARDL Bounds test

The last crisis indicator is derived from ARDL Bounds test. This method tests whether a long-run relationship exists between a dependent variable (real exchange rate) and a set of explicative variables (international reserves and nominal interest rate), when it is not known with certainty whether the underlying are trend- or first-difference stationary. The Bounds tests are based on standard F- and t-statistics. The asymptotic distributions of these statistics are non-standard under the null hypothesis that there exists no level relationship, irrespective of whether the variables are I(0) or I(1). Two sets of asymptotic critical values are provided: one when all variables are I(1) and the other if they are all I(0). These two sets of critical values provide a band covering all possible classifications of the variables into I(0), I(1) or mutually cointegrated (Pesaran et al., 2001).[7] The model we test through ARDL procedure is as follows:

$$\Delta RER = \alpha_0 + \sum_{i=1}^{p} \alpha_1 \Delta RER_{t-i} + \sum_{i=1}^{p} \alpha_2 RES_{t-i} + \sum_{i=1}^{p} \alpha_3 \Delta NIR_{t-i} + \lambda_1 RER_{t-1} + \lambda_2 RES_{t-1} + \lambda_3 NIR_{t-1} + \varepsilon_t \tag{13}$$

where the expression from λ_1 to λ_3 illustrates the long-run relationship between the variables, while the expression from α_1 to α_3 with the summation signs corresponds to the short-run dynamics of variables. On the other hand, α_0 represents constant and ε_t is Gaussian noise.

The complete results for short- and long-run dynamics through ARDL Bounds testing are obtained through several steps. In the first step, Eq. (13) is estimated through OLS method and F-test is conducted to test the existence of a long-run relationship among the variables. The null hypothesis is Ho: $\lambda_1 = \lambda_2 = \lambda_3 = 0$. This means the non-existence of a long-run relationship. The calculated F-statistics value is compared with upper and lower critical values which are given by Pesaran et al. (2001). If calculated F-value exceeds the upper critical value, then null hypothesis of no-cointegration is rejected irrespective of whether the variable are I(0) or I(1).

As shown in Table 2 where ARDL Bounds test results are given, F-statistic exceeds the upper critical value. This means that there is a long-run relationship between real exchange rate, reserves and nominal interest rate; it thus confirms the speculative pressure approach. This relationship also constitutes our twelfth crisis index. We set then four values of threshold above the index mean to obtain crisis dates for the Turkish economy.

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[7] See Pesaran et al. (2001) for further details on ARDL Bounds test.

Table 2. ARDL Bounds test results

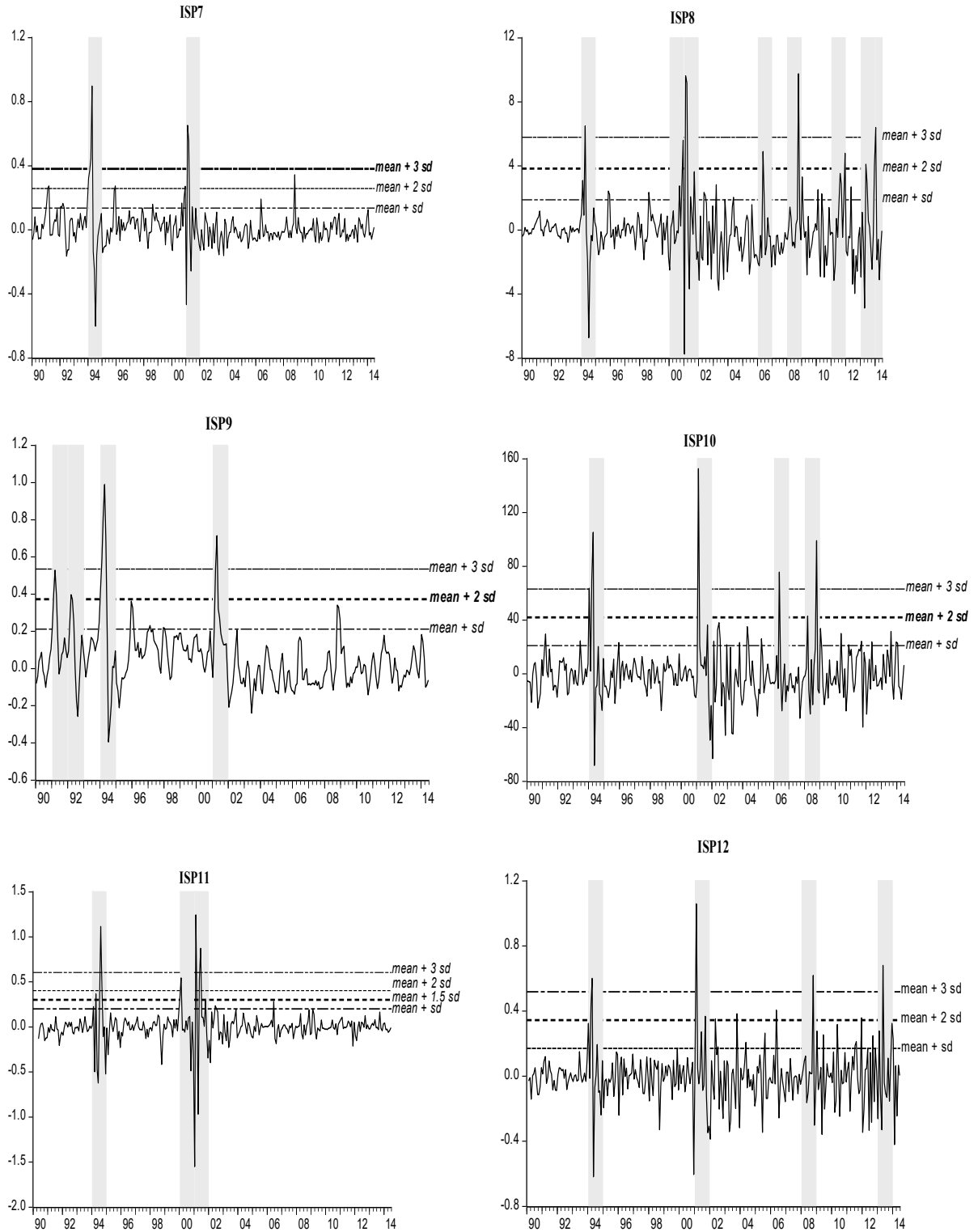|	F statistic	5% C.V.		10% C.V.	
		I(0)	I(1)	I(0)	I(1)
F(RER| RES, NIR)	7.54	3.88	4.61	3.38	4.02
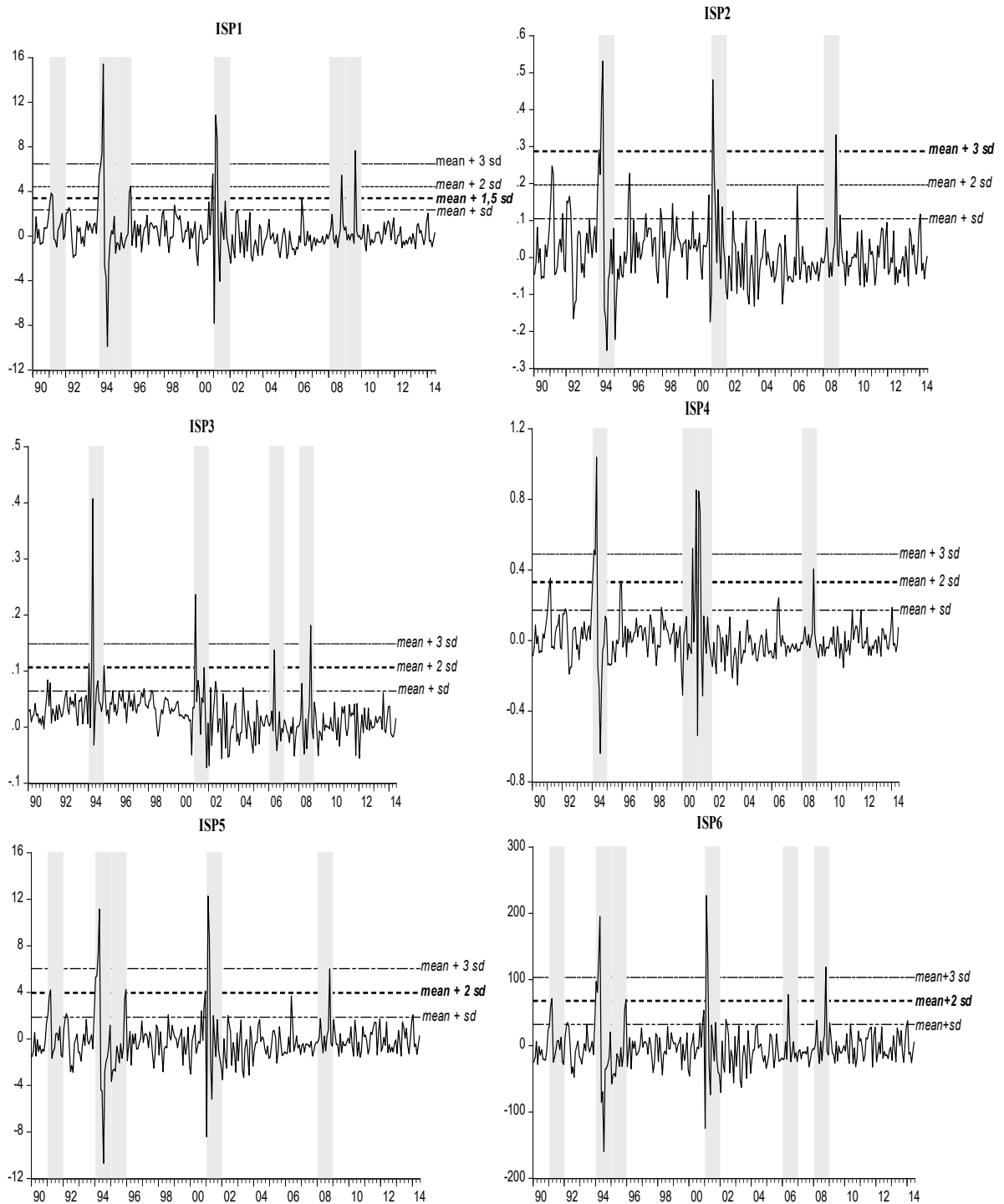
Figure 1. Currency crisis indicators at different thresholds

Note: The bold line represents the original value of threshold used in those original papers.

3.3 Time Series Properties of ISPs

Table 3 presents descriptive statistics of the ISP indexes. As seen in Table, ISP6 and ISP10 have the highest mean, median, and standard deviation. The common point of these two crisis indicators is the use of real exchange rate, contrary to others that employ the nominal exchange rate. On the other hand, ISP2, ISP3, ISP4, ISP7, ISP11 and ISP12 own the lowest mean, median, and standard deviation. The common point of these indicators is the weighting procedure: as mentioned above, the ISP3 sets arbitrary (0.75 and 0.25) weights to the index components, while the ISP4 use standardised weights. On the other hand, the weight of each component in the ISP2 and ISP7 is normalized by the sample standard deviation of the nominal exchange rate. Besides, in the Cointegration and ARDL methods, the weighing procedure is realized through long-run estimations of the model. All these observations show that the weighting procedure affects the volatility of the index, but it does not have any significant impact on dates of currency crises (see Figure 1). This result also confirms the findings of Eichengreen et al. (1996) and Corsetti et al. (2001).

Table 3. Descriptive statistics of ISPs

ISP	Mean	Median	Standard deviation	Skewness	Kurtosis	Jarque-Bera
ISP1	0.293	0.022	2.064	1.954	18.225	3026.420
ISP2	0.014	0.002	0.091	1.686	9.999	739.442
ISP3	0.022	0.021	0.042	3.137	27.974	8122.514
ISP4	0.013	-0.010	0.159	2.215	15.782	2241.773
ISP5	-0.217	-0.402	2.084	1.269	13.634	1464.020
ISP6	-3.285	-6.864	35.511	1.696	14.155	1665.190
ISP7	0.013	-0.001	0.123	1.824	17.909	2886.142
ISP8	-0.076	-0.177	1.958	1.234	9.625	612.265
ISP9	0.050	0.023	0.162	1.766	9.561	680.024
ISP10	-0.161	-2.504	21.110	2.162	15.990	2296.081
ISP11	-0.003	-0.006	0.203	-0.393	24.513	5599.870
ISP12	-0.003	-0.007	0.174	0.928	9.453	550.3784

Table 4. ADF and PP tests of ISPs

Variable	ADF unit root test		PP unit root test	
	T-statistic (Level)	T-statistic (First difference)	T-statistic (Level)	T-statistic (First difference)
ISP1	-13.03238***	-12.28307***	-12.79489***	-81.37454***
ISP2	-11.89702***	-12.32314***	-11.47648***	-70.69109***
ISP3	-15.76166***	-12.30556***	-15.84364***	-139.7515***
ISP4	-13.28889***	-12.04237***	-13.27073***	-84.50486***
ISP5	-13.23902***	-12.24678***	-12.93136***	-78.84778***
ISP6	-13.47094***	-12.26597***	-13.23278***	-87.79695***
ISP7	-12.70956***	-12.39787***	-12.17247***	-76.16100***
ISP8	-15.00709***	-12.48979***	-14.94613***	-99.56436***
ISP9	-5.560618***	-9.734588***	-4.341345***	-18.45469***
ISP10	-15.34025***	-12.47011***	-15.25206***	-111.6730***
ISP11	-9.863131***	-10.55662***	-25.37999***	-112.6092***
ISP12	-16.20365***	-12.28090***	-22.10558***	-140.7688***

Note: *** represent statistical significance at the 1% level.

The skewness and kurtosis results in Table 3 indicate that ISP distributions have fat tails. All the ISP series, except for the ISP11, are skewed to the right with kurtosis coefficients exceeding the value of 3 found for normal distributions. Jarque-Bera statistics give evidence that the null hypothesis of normality is rejected in all cases. This suggests that the crisis indicators follow a normal distribution, and the threshold method in terms of a number of standard deviations above the mean is validated. Table 4 shows the results of two types of unit-root tests (ADF

and PP) for the crisis indicators. All the indicators are stationary at level and first difference, since the unit root null hypothesis of both tests is rejected at the 1% level.

3.4 Comparison

This section compares the performance of the ISP indicators in identifying crisis episodes in the Turkish economy in the 1990-2014 period. First, all the indicators perform similarly: 1994 and 2001 crises are detected by every index even at the highest threshold; while 1991, 2006 and 2008 are generally identified as crisis episodes at lower thresholds.

Second, the threshold value significantly matters. We observe that as expected the higher the value of the threshold, the lower the number of crises. For example, at threshold of 1 standard deviation above the index mean, the number of identified crisis episodes exceeds easily 20, while at threshold of 3 standard deviations, the number of identified crisis episodes is 5 on average. These results show that the main problem with the existing literature on crisis dating is not the definition of currency crisis but the set of different threshold levels. Hence, it is crucial to make sensitivity tests by utilizing different values of threshold. Of course, the only important thing here is not the number of identified crises; the dates of identified crises should also match to the dates of actual or past crisis episodes. In this sense, one can observe through Figure 1 and Table 6 that all the indicators perform quite well.

Table 5. Number of crisis: Sensitivity to definition of ISP and choice of threshold

Variable	Threshold	Number of Crisis	Variable	Threshold	Number of Crisis
ISP1	1	21	ISP7	1	23
	1.5	13		1.5	13
	2	10		2	11
	3	5		3	4
ISP2	1	31	ISP8	1	31
	1.5	18		1.5	17
	2	10		2	10
	3	5		3	5
ISP3	1	26	ISP9	1	27
	1.5	8		1.5	17
	2	6		2	11
	3	3		3	6
ISP4	1	23	ISP10	1	29
	1.5	14		1.5	13
	2	10		2	7
	3	6		3	6
ISP5	1	22	ISP11	1	14
	1.5	13		1.5	8
	2	10		2	6
	3	4		3	3
ISP6	1	23	ISP12	1	27
	1.5	13		1.5	17
	2	10		2	10
	3	5		3	4

Third, as we do not use any exclusion window, at lower values of threshold same crisis episodes are signalled as different crisis. The most significant example of this issue is the identification of the April 1994 crisis (see Table 6). Thus, the number of identified currency crises at different values of threshold should be interpreted cautiously, by taking into account this exclusion window issue.

Fourth, the exchange rate whether it is nominal or real seems us the most important variable to illustrate currency crises. Therefore, including reserves and/or interest rates have little impact on the crisis index, thus in identifying crisis episodes. However, as indicated by the Cointegration and ARDL tests, there exists a long-run relationship between the exchange rate, reserves and interest rates. This means that including reserves and/or interest rates is not useless.

Fifth, using reserves differential between Turkey and the reference country (the US) means nothing (see ISP1 in Figure 1), but using interest rate differential (see ISP8 in Figure 1) makes sense in a country which is closely tied to international financial markets. ISP8 is the only indicator that shows the period of high fluctuations in financial markets during the period May-December 2013. This period actually corresponds to corruption investigations and 'Gezi' manifestations across the country.

Table 6. Crisis dates by crisis indicators at threshold of 2 standard deviations

ISP1		ISP2		ISP3	ISP4		ISP5		ISP6	
94Jan.	00Dec	91Mar.	00Dec.	94Jan.	91Mar.	00Sep.	91Mar.	95Dec.	91Mar.	95Dec.
94Feb.	01Feb.	94Jan.	01Jan.	94Mar.	94Jan.	00Dec.	94Jan.	00Dec.	94Jan.	01Feb.
94Mar.	01Mar	94Mar.	01Feb.	94Apr.	94Feb.	01Feb.	94Feb.	01Feb.	94Feb.	01Mar.
94Apr.	08Oct.	94Apr.	06May	01Feb.	94Mar.	01Mar.	94Mar.	01Mar.	94Mar.	06May
95Dec.	09Aug.	95Dec.	08Oct.	06May	94Apr.	08Oct.	94Apr.	08Oct.	94Apr.	08Oct.
				08Oct.						

ISP7		ISP8		ISP9		ISP10	ISP11	ISP12	
91Mar.	00Dec.	94Apr.	08Oct.	91Feb.	94Mar.	94Jan.	94Apr.	94Mar.	03Oct.
94Jan.	01Feb.	00Dec.	11Dec.	91Mar.	94Apr.	94Mar.	94Aug.	94Apr.	06May
94Mar.	01Mar.	01Feb.	13May	91Apr.	94May	94Apr.	00Jan.	01Feb.	08Oct.
94Apr.	08Oct.	01Mar.	13Dec.	92Mar.	01Mar.	01Feb.	01Feb.	01Sep.	11Dec.
95Nov.	08Nov.	06May	14Jan.	94Jan.	01Apr.	06May	01May	02May	13May
95Dec.				94Feb.		08Mar.	01June		
						08Oct.			

Another important point that should be underlined how the distribution of the crises over time is. As seen in Figure 2 the crisis episodes are not randomly distributed. As expected, the Figure shows a clustering of crises in 1994, 2001, 2008, and also less significantly in 1991 and 2006. This clustering of crisis episodes is consistent with the crises occurred in the Turkish economy.
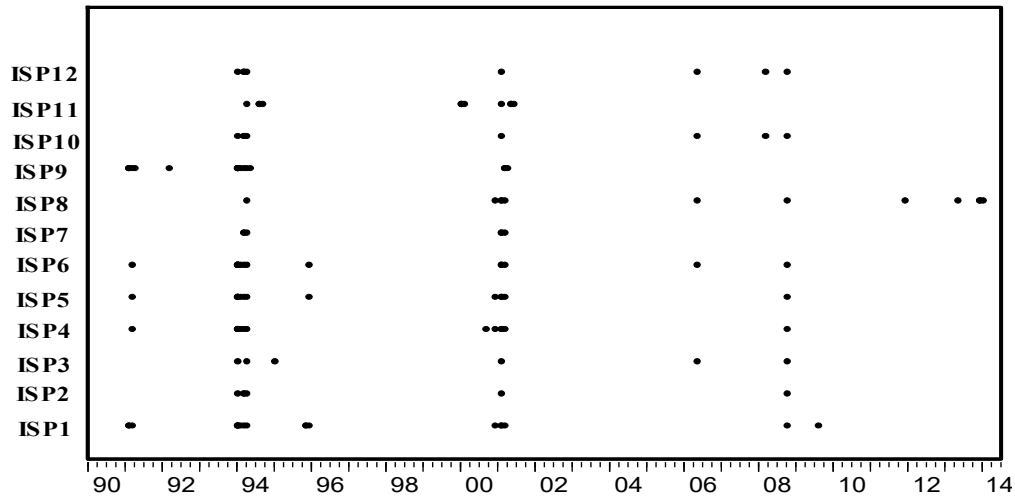
Figure 2. Distribution of currency crises over time

CONCLUSION

This paper reproduced, for the Turkish economy, a broad set of the ten most representative indicators from the existing literature and constructed two new crisis indicators which are derived from Cointegration and ARDL Bounds tests.

As we already underlined, our different currency crisis indicators performed well and similarly in identifying correctly the severe currency crises of 1994, 2001 and 2008 in Turkey. Moreover, our results indicate that the weights assigned to the index components matter little, but threshold values matter significantly in dating currency crises. Besides, using exclusion window leads to information loss on the duration of currency crises, but may be useful in detecting the correct dates of crises, particularly if one opts for a low value of threshold. In addition, the nominal or real exchange rate seems the most significant and reliable variable to illustrate currency crises. However, as indicated by the Cointegration and ARDL tests, there exists a long-run relationship between the exchange rate, reserves, and interest rates. This means that reserves and/or interest rates have an impact on crisis index, and also on crisis dates.

This research suggests no indicator could perfectly track crisis episodes, thus using different crisis definitions and different values of threshold in order to assess the robustness of the crisis dating schemes is important. The results of this research are country-specific, but this does not shadow the originality of this paper and its significant contributions to the existing literature.

Further research derived from these conclusions consists of testing whether these (small) differences in crisis periods identification influence how the early warning systems work.

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