Dynamic Linkages of Current Account Deficits and Unemployment: Evidence from Turkey

Mustafa Özer
Department of Economics
Eskişehir Anadolu University

and

A. Erinç Yeldan
Department of Economics
Yaşar University

Abstract

In this paper, we empirically test the causal relationship between current account deficits and unemployment in Turkey over 2000Q1–2012Q1. Using Johansen co-integration and Granger-causality analyses based on a corresponding vector error correction model, we studied many alternative specifications of the nexus between unemployment and external deficits in an open macroeconomy environment. Our results reveal the presence of unidirectional causality running from current account deficits to unemployment. Furthermore, based on the impulse response variance decomposition analysis, we find that unemployment explains little variation of current account deficits, although current account deficits explain a substantial fraction of the variation in unemployment. We interpret these findings as evidence of the structural sources of unemployment being embedded under the deepening external fragility of the Turkish economy over the 2000s.

Keywords: Turkey, Current account deficit; Unemployment; Unit root; Cointegration; Granger causality; VECM

1. Introduction

During the 2000s, despite rapid growth and a significant surge in exports, Turkish economy could not generate jobs at the desired rate. Open unemployment rate which stood at 6.5% in 2000, has jumped to 10.3% in 2002 in the aftermath of the February 2001 financial crisis. Since then the Turkish gross domestic product has increased by a cumulative 30% in real terms. Yet, employment generation capacity of this rapid growth had been dismal, and the open unemployment rate could not be brought down below 9% by the end-of 2007, just before the eruption of the current global economic crisis. Despite rapid expansion of production in many sectors, civilian employment increased sluggishly at best, and labor participation remained below its levels as observed during the 1990s.

A further key distinguishing feature of the Turkish economy over the 2000s was the eruption of the current account deficits in almost a structurally permanent manner. Traditionally Turkey used to display a fair balance in its current account. However, starting 2003 annualized current account deficit, as a ratio to the gross domestic product, increased to the 3 – 4% band, and then jumped above 6% after 2006 to reach a record high 9.7% in 2011.

Our working hypothesis in this paper is that the meager job creation in Turkey over 2000s is the direct symptom of a speculative-led growth environment (*a la* Grabel, 1995) together with an excessively open and unregulated capital account in the age of relatively cheap and abundant global finance. Accordingly, with the available bonanza of relatively cheap external credit, Turkey could have financed its imports via rapid accumulation of external debt. Substitution of imports for domestic production led to lower value added production at home. Thus, the problem of poor job performance and the fragility embedded in the increase of the current account deficits were, in fact, manifestations of the same adjustment mechanism under a speculative finance-led growth path.

We study this hypothesis utilizing time series econometrics based on Johansen co-integration and Granger causality techniques. Focusing on quarterly data over 2000 to 2012, we investigate the relationships of co-integration and causality between unemployment and external balance for the Turkish economy. Our findings reveal that current account deficits explain a substantial fraction of the variation in unemployment and suggest the presence of strong unidirectional causality. The plan of the paper is as follows: In the next section we provide an overview of the macroeconomics of external deficits and employment performance of the Turkish economy over the 2000s. We introduce our econometric methodology in section 3, and implement a series of econometric tests in section 4. Section 5 concludes and suggests policy implications for future research.

2. Patterns of External Deficits and Unemployment in Turkey over the 2000s[1]

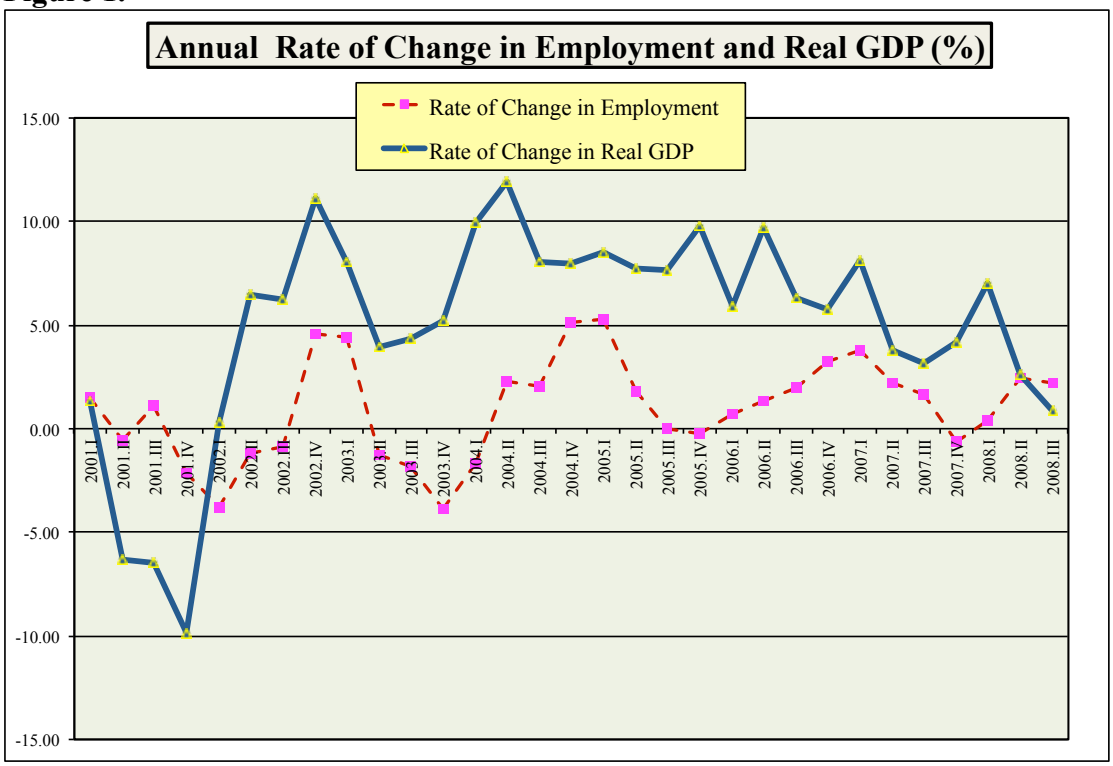
As a newly emerging market economy Turkey had been subject to the patterns of the global business cycle over the 2000s. During the 1990s, the economy suffered from a high inflationary environment with unsustainable fiscal deficits. The unfavourable macro economic setting culminated into two severe financial crises in 1994 and 2001. Both of these were driven by speculative attacks of foreign finance capital driven by the unsustainable rates of return under conditions of deep fiscal and external fragility.

In what follows, the post-2001 crisis adjustments came at a very unique conjuncture of the global economy. First of all, growth, while rapid, showed quite peculiar characteristics. It was mainly driven by a massive inflow of foreign finance capital which, in turn, was lured by significantly high rates of interest offered domestically; hence, it was *speculative-led* in nature (Grabel, 1995). The main mechanism has been that the high rates of interest prevailing in the Turkish asset markets attracted short term finance capital, and in return, the relative abundance of foreign exchange led to overvaluation of the *Lira*. Cheapened foreign exchange costs led to an import boom both in consumption and investment goods. The overvaluation of the *Lira*, together with the greedy expectations of the arbitrageurs in an era of rampant financial glut in the global finance markets, led to a severe rise in its foreign deficit, and hence, in external indebtedness.

A further characteristic of the post-2001 era was Turkey's *poor job creation* pattern. Rapid rates of growth were accompanied by high rates of unemployment and low participation rates. The rate of total unemployment rose to above 10% after the 2001 crisis, and despite rapid growth, has not come down to its pre-crisis levels. In fact, the most relevant observation from this history is that during the 2000s, despite rapid growth and a significant surge in exports, Turkish economy could not generate jobs at the desired rate. To make this assessment clearer we plot the quarterly growth rates in real gross domestic product in Figure 1, and contrast the *y-o-y* annualized rates of change in labour employment. In order to make comparisons meaningful, the changes in labor employment is calculated relative to the same quarter of the previous year.

[1] Parts of this section draw on the workshop on "Patterns of Growth and Employment in Turkey" organized by the International Labour Organization, Turkey Office, November 2012.

Figure 1.



Source: Derivations from Turkstat Household Labour Force Statistics and National Income Accounts

The figure discloses that **over 27 quarters** of data points between 2002.Q1 and 2008.QIII (the date of the contamination of the global crisis), the average rate of growth in real GDP had been 6.5%. In contrast, the rate of change of employment averaged *only 0.8%* over the same period. Over the twenty seven quarters portrayed in the figure, GDP growth was *positive* in all periods. Yet, labor employment growth was *negative* in 10 of those 27 quarters. Another reflection of this phenomenon was the significantly *low elasticity of employment*; that is percentage gain in employment due to percentage changes in GDP growth had been relatively low (see table 1 below).

Compared over broad period averages, employment generation capacity of the domestic economy seems to have been relatively poor in the post-2000s. There had been labor shedding in agriculture, while the non-agricultural sectors had significantly lower employment elasticities. All of these phenomena had been succinctly phrased as *jobless growth* for Turkey. (see, *e.g.* Telli, Voyvoda, Yeldan, 2006; Voyvoda and Taymaz, 2009).

Table 1

Output Elasticities of Employment By Sectors (Annual averages)

	1989-2008	1989-2000	2002-2008
Total	0.25	0.39	0.14
Agriculture	-1.19	-0.42	-1.66
Non-Agricultural Sectors	0.54	0.68	0.48
Industry	0.43	0.49	0.39
Services	0.55	0.76	0.47

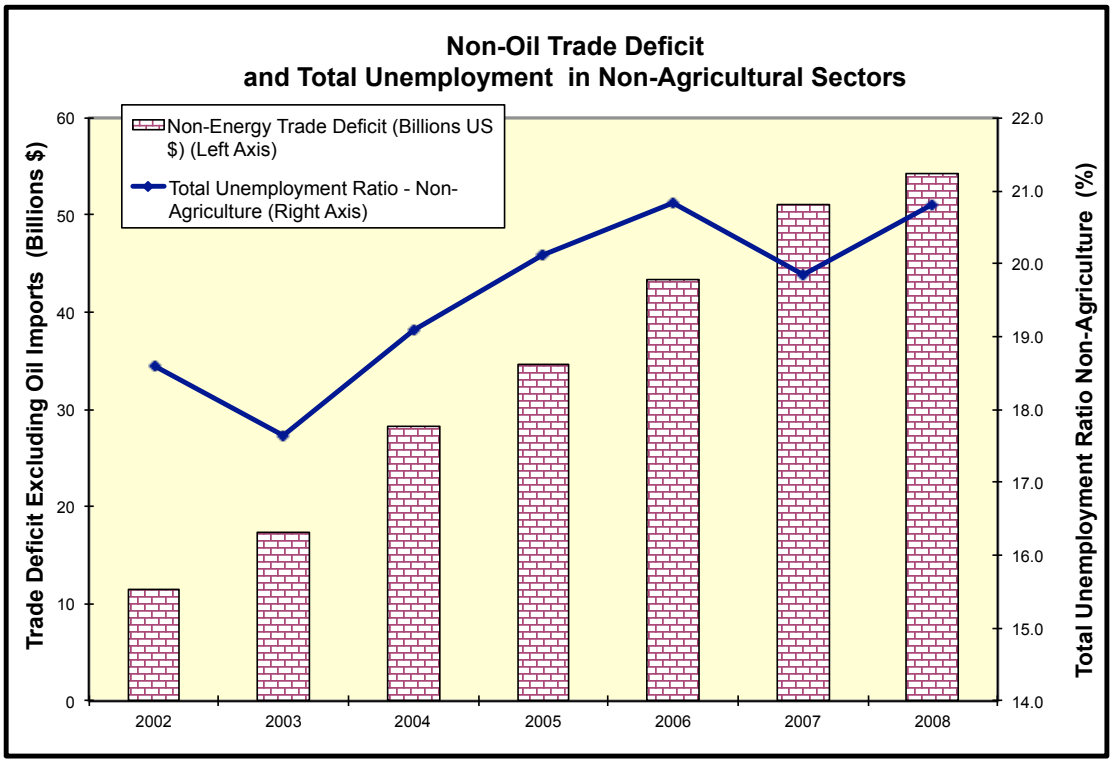
Source: Author's calculations based on Turkstat and SPO data

Based on these observations a *macroeconomics driven* hypothesis can be formulated to argue that the jobless growth problem is a direct symptom of the current macroeconomic program as implemented in Turkey together with an excessively open capital account and widespread financial speculation. According to this line of thought, due to virtually unregulated capital account, and given the relatively high real rates of interest prevalent in the Turkish financial markets, Turkey is observed to receive massive inflows of short term finance capital. As a result, the domestic currency, *TL*, appreciates and Turkey suffers from a widening current account deficit. Appreciated currency brings forth a surge in imports together with a contraction of labour intensive, traditional export industries such as textiles, clothing, and food processing. This leads to contraction of formal jobs and increased informalization of economic activities (see Yeldan (2006, 2011), Telli, Voyvoda and Yeldan, 2006).

The structural overvaluation of the TL, not surprisingly, manifests itself in ever-expanding deficits on the commodity trade and current account balances. In what follows, starting in 2003 Turkey has witnessed expanding current account deficits, with the figure in 2011 reaching a record-breaking magnitude of $78.1 billion, or 9.7% as a ratio to the aggregate GDP. In appreciation of this figure, it has to be noted that Turkey traditionally has never been a current account deficit-prone economy. Over the last two decades (80's and 90's) the average of the current account balance hovered around plus and minus 1.5-2.0%, with deficits exceeding 3% typically signaling significant currency adjustments.

The close relationship between meager job creation and the foreign deficits are portrayed succinctly in Figure 2. Here, in order to isolate for the effect of non-energy imports, the size of non-oil trade deficit is portrayed in reference to the right hand-axis. Due to the presence of high seasonality and structural factors, the rural economy is also taken as exogenous to the Figure 2. Thereby, we follow the close relationship of the *non-oil trade deficit* together with the *non-agricultural unemployment*.

Figure 2

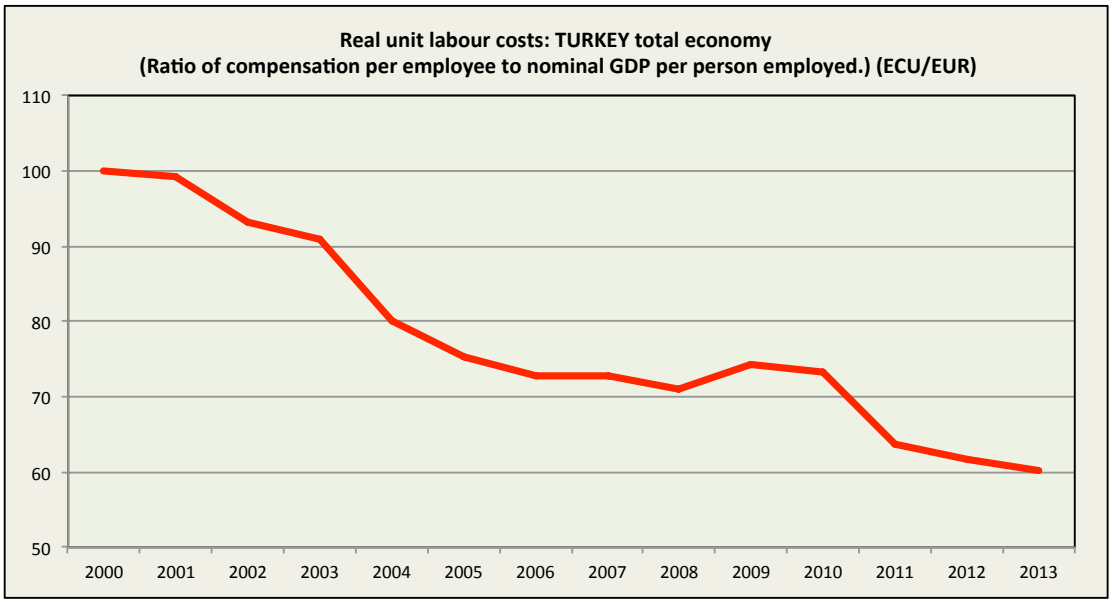


Source: Turkstat Household Labour Force Statistics

The portrayal of the rising non-agricultural unemployment along with an expanding (non-oil) trade deficit is no surprise to students of development economics. As Turkey consumed more and more of value added produced *abroad*, and found it profitable to do so with an appreciated currency financed by speculative financial inflows, external deficit widened and foreign debt accumulated. The costs of this *speculative-led growth*, however, were realized as loss in jobs, deepening informalization, and decline of real wage income.

The evolution of the real unit labor wage costs as calculated by the European Commission reveals these trends succinctly (see Figure 3). Weighted by the productivity indexes, the fall in real unit labor costs indicates that the loss in export competitiveness dur to currency appreciation could have been overcome by depressing the real wage costs against productivity gains. In this way, firms could have sustained a competitive edge in the global commodity markets.

Figure 3



Source: European Commission Economic and Financial Affairs, AMECO data base

All these observations leave us with the working hypothesis that the persistent unemployment problem in Turkey over the 2000s has strong structural features rooted in the externally fragile macro economic environment. It is this issue that we now turn in more formal terms.

3. Methodology

This section highlights the methodologies that this paper uses to explore the dynamic linkages between current account deficits (CAD) and unemployment in Turkey over the 2000Q1–2012Q1 period. As pointed out in (Cheng, 1999, p. 912), we prefer to use Granger causality test method to examine possible causal relationships between CAD and unemployment in Turkey in light of Monte Carlo evidences as provided by Guilkey and Salami (1982) and by Geweke et al. (1983).

To carry out the Granger causality test, we first investigate the order of integration of the series using different unit root tests and then test existence of cointegration between series by employing Johansen cointegration test to confirm that the Granger causality tests will not produce any spurious results (AuYong et al., 2004, p. 481). Secondly, we try to identify the short-run and long-run causality between two variables using VECM framework. Finally, based on the impulse responses and variance decomposition, we try to analyze the dynamic relations between the two series.

Since Granger causality test requires the determining the order of integration series, we first examined stochastic properties of two series by applying Augmented Dickey-Fuller (ADF) test (Dickey-Fuller, 1979), Phillips Perron (PP) test (1988), Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test (Kwiatkowski et al., 1992), and Zivot and Andrews (ZA) test (1992). We provide the specifics of this analysis in Appendix 1.

3.1. Granger Causality Tests

The Granger (1969) causality test is often conducted in the context of a vector autoregression (VAR). It is designed to detect direction of the possible causal relationship between two time series by examining a correlation between the current value of one variable and past values of another variable. According to Granger (1969), X Granger causes Y, if current value of Y can be predicted better by taking into account of past values of X than by not doing so, provided that all other past information in the information set is used. To carry out the Granger causality test, degree of the integration of the variables should be known to avoid the spurious inferences.

If the series CAD and UNEMP are individually integrated of order one, but not co-integrated, then to test the Granger causality, a VAR model in first differences should be used, since taking first differences make the series stationary series. To this end we specify the following equations:

$$\Delta unemp_t = \alpha_1 + \sum_{i=1}^{p} \beta_i \Delta cad_{t-i} + \sum_{j=1}^{q} \delta_j \Delta unemp_{t-j} + \varepsilon_{1t} \tag{1}$$

$$\Delta cad_t = \alpha_2 + \sum_{j=1}^{q} \phi_j \Delta unemp_{t-j} + \sum_{i=1}^{p} \gamma_i \Delta cad_{t-i} + \varepsilon_{2t} \tag{2}$$

where $\Delta unemp$ and Δcad are the first differences in these variables that capture their short-run disturbances, ε_{1t} and ε_{2t} are the serially uncorrelated errors and ect_{t-1} is the lagged error correction term, which is derived from the long-run cointegration relationship and measures the magnitude of the past disequilibrium.

Based on the eq. (1), we can test whether current account deficits Granger-causes unemployment by testing statistical joint significance of the coefficients of the lagged CAD by a joint-F test. The existence of co-integration between the two series implies that the existence of Granger causality, but it does not indicate the direction of the causal relationship (Granger, 1988). Therefore, vector error correction model (VECM) should be employed to detect the direction of the causality. As stated in (Belloumi, 2009, p.2749), the VECM allows to distinguish between long-run and short-run causality between variables and can identify the sources of the causation that cannot be determined by the traditional Granger causality test. As mentioned in Granger (1988), the dynamic Granger causality can be obtained from the VECM derived from co-integrating relationship. The VECM for the CAD and unemployment can be written as follows:

$$\Delta unemp_t = \alpha_1 + \sum_{i=1}^{p} \beta_i \Delta cad_{t-i} + \sum_{j=1}^{q} \delta_j \Delta unemp_{t-j} + \varphi_1 ect_{1t-1} + \varepsilon_{1t} \tag{3}$$

$$\Delta cad_t = \alpha_2 + \sum_{j=1}^{q} \phi_j \Delta unemp_{t-j} + \sum_{i=1}^{p} \gamma_i \Delta cad_{t-i} + \varphi_2 ect_{2t-1} \varepsilon_{2t} \tag{4}$$

The coefficients, φ_1 and φ_2 of ect_{t-1}., measure the error correction mechanism that derives the variables back to their long-run equilibrium relationship.

Using Eqs. (3) and (4), we can have the following different cases of causal relations (short-run Granger causality) based on the Wald χ^2-test; (i) current account deficits Granger-cause unemployment only when lagged values of Δcad in Eq. (3) may be statistically different from zero while values of $\Delta unemp$ are not in Eq. (4). The joint significance of the coefficients of lagged values of CAD variable indicates that the unemployment responds to short-run shocks to the stochastic environment. (ii) Unemployment Granger-cause current account deficits only when lagged values of $\Delta unemp$ in Eq. (4) may be statistically different from zero while values of Δcad are not different from zero in Eq. (3); (iii) bidirectional causality occurs when both the lagged values of Δcad and $\Delta unemp$ in Eqs. (3) and (4) are significantly different from zero and (iv) there is no causal relation between current account deficits and unemployment when both the lagged values of Δcad and $\Delta unemp$ in Eqs. (3) and (4) are significantly not different from zero. In this case, we can conclude that the variables are independently moving on their paths without influencing each other.

We can detect presence of long-run causality by testing the significance of coefficient of the error correction term (*ect*), which is the speed of adjustment, by applying separate t-tests on the speed of adjustment coefficients (φ_1 and φ_2). The significance of the speed of adjustment term indicates that the long-run equilibrium relationship is directly driving the dependent variable. As pointed out in (Abbas and Choudhury, 2012, p.7), as disequilibrium error term is integrated of order of zero, which is a stationary variable, there will be some adjustment process preventing the errors in the long-run becoming larger. Also first differenced Error Correction Model (ECM) eliminates the trends from the variables in the model resolving the problem of spurious regression. Also, reintroduction of lagged error correction term (ect_{t-1} in Eqs. (3) and (4)) captures the long-run information lost through differencing.

3-2. Impulse Response Analysis

As is well known, impulse response function (IRF) allows us to capture the dynamic behavior as it traces the effect of an exogenous shock to a variable on current and future values of another variable in the VAR system. It also takes into account that variables have a common component (Glass, 2009, p. 31). With the use of this function, one can have the ability to describe the dynamic interplay among the variables and observes the adjustment speed of variables in the model (Yu et al., 2008, p. 59). Through the dynamic interactions among variables, the disturbance in error term of a certain variable in period t will cause series of changes to all variables in the model after period t. Clearly, many other explainable factors are the reasons for possible disturbance of a variable that stands against such shocks.

Even though the VECM Granger causality approach provides a powerful means for determining both short-and long-run Granger causality tests, especially to capture the direction of the causality between series, it does not tell how the series respond where there is a shock in one of the variables within the system. In other words, even though the VECM Granger causality approach allows determining the direction of Granger causality, it does not tell the sign of the causality. To determine the sign of the causality, a number of prior studies uses the sum of the coefficients but, as argued in (Le and Chang, 2013, p.85), this approach may produce misleading results as there are all of the dynamic effects between the Eqs. (6) and (7) that have to be taken into account.

As is suggested in (Le and Chang, 2013), to capture the sign of the Granger causality, one has to look at the sign of the impulse responses for all periods. If the response function is positive for all periods, fading away to zero, this should be taken as an indication of positive causality. But on the other hand, it is positive, then negative, and then dampens down; it may be interpreted as a sign of absence of a clear-cut sign of causality. In this case, it could be said that the sign of causality depends on the time horizon.

3-3. Variance Decomposition Analysis

As pointed out in (Akinlo, 2009, , p. 686-687) and mentioned in (Shahbaz, 2012) also, by employing the VECM Granger causality test, one can only be able to test the causality among variables within the sample period. In other words, with the VECM Granger causality test, we cannot detect the relative strength of causal relationship beyond selected sample period. This result usually is considered as limitation of such test and it weakens the reliability of VECM Granger causality test results. Therefore, to capture the out of sample causality, the use of variance decomposition analysis is recommended. By portioning the variance of the forecast error of a certain variable, say unemployment, into proportions attributable to shocks in each variable, such as current account deficits, in the system including its own, VDCs might indicate Granger causality beyond the sample period. Also, with the VDCs analysis, one can capture the magnitude of the predicted error variance for a series accounted for by innovations from each of the independent variable over different time-periods beyond the selected sample period.

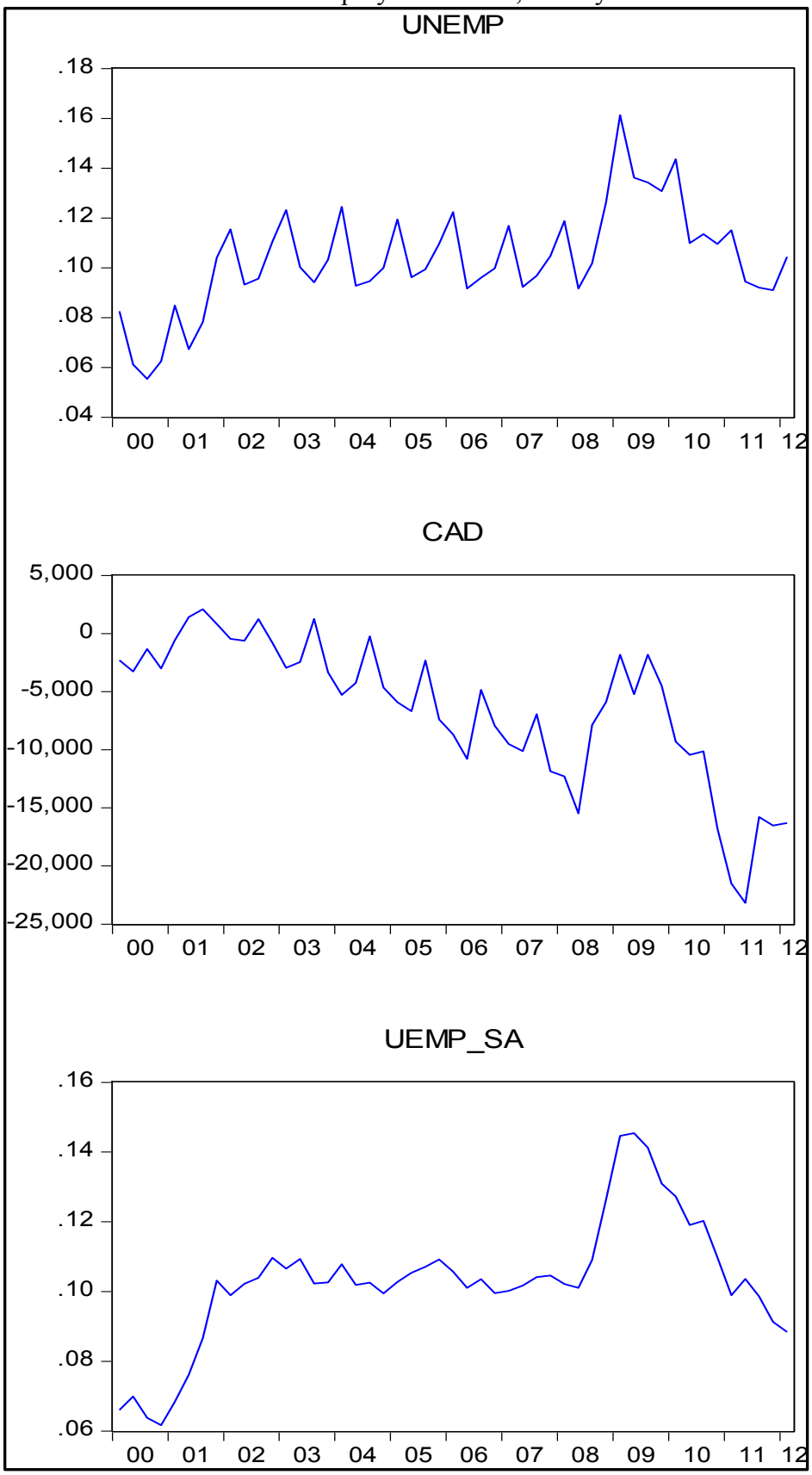
4. Empirical Results

In this section, we interpret the results of the econometric techniques that we highlighted above and discuss their implications.

4.1. Data

In this study, we used quarterly data for Turkey from 2001Q1 to 2012Q2, as tabulated by the Central Bank of the Rep. of Turkey and Turkish Statistical Institute, Turkstat. Figure 4 displays the time series plots of all variables used in the study. Since unemployment exhibits clear seasonality, we used Tramo/Seats method to remove the seasonal component in unemployment series. Also time series plots of all variables shows a structural break following the 2008/09 crisis.

Figure 4. Unemployment, current account deficits, and seasonally adjusted unemployment series, Turkey.



In order to identify the order of the integration of the variables in our study, we carry out unit root and stationarity tests of ADF, PP and KPSS, without a trend. These tests are reported in Appendix 2.

4-2. Granger causality test results

Since the CAD and UNEMP variables are co-integrated (see Appendix 2), we set up a VECM for examining the short-run and long-run causalities and determine the direction of causality. In the VECM, we regressed the first difference of each endogenous variable on a ten period lag of the co-integrating equation. To examine the short-run causality, we implemented non-causality standard Wald chi-square test to compute the joint significance of the lagged differences of the explanatory variables in the short-run. For the long-run causality, we applied the t-test on the coefficient of error correction term. Table 2 shows the results of VECM Granger causality tests.

Table 2

Results of VECM Granger causality tests.

Long-run		**Short-run**	
Null hypothesis	t-stat.	Null hypothesis	χ^2-stat.
ΔCAD does not Granger cause $\Delta UNEMP$	-3.02**	ΔCAD does not Granger cause $\Delta UNEMP$	23.22**
$\Delta UNEMP$ does not Granger cause ΔCAD	0.55	$\Delta UNEMP$ does not Granger cause ΔCAD	16.20
*indicates the statistical significance at 5% level of significance.			

Based on the results of the VECM Granger causality tests in Table 2, we found that there is a unidirectional causality running from current account deficits to unemployment at the 5% significance level in the short-run, but the converse is not true. Moreover, the results of the long-run Granger causality test of the ECT confirm that the ECT coefficient of unemployment is negative and statistically significant at the 5% significance level and the ECT coefficient of the CAD is positive but insignificant.

The results of the long-run causality tests imply that current account deficits is a weakly exogenous variable, but unemployment is not; therefore indicating the presence of unidirectional causality running from CAD to UNEMP in the long-run. The value of the ECT coefficient of unemployment is approximately -0.47. This implies that the adjustment coefficient (speed of adjustment) is 47% in the equation indicating that the corrections to the short-run disequilibrium will be 47% per year.

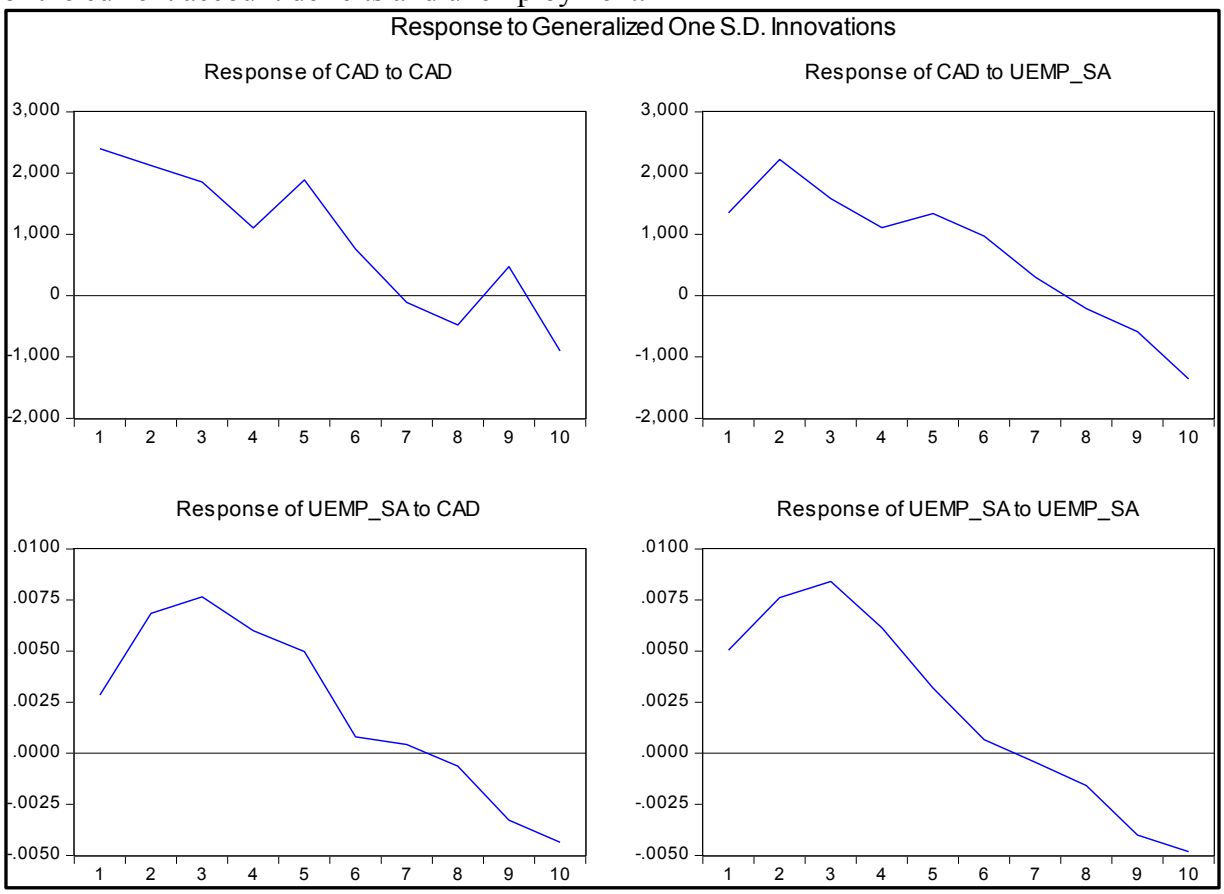
4-3. Impulse responses

To further investigate the dynamic response between current account deficits and unemployment, particularly to get some idea about the sign of the unidirectional causality detected in previous section running from current account deficits to unemployment, we also

calculate the impulse response of the VECM based on the generalized impulse responses, since it is not subject to orthogonality critique. Figure 5. displays the impulse responses of VECM for the variables where Granger causality was detected.

Figure 5

Impulse responses to generalized one standard deviation innovations according to the VECM of the current account deficits and unemployment.



Impulse response of unemployment to a shock in current account deficits has a positive sign most of the time (shocks in current account deficits has positive effect on unemployment about two years, and then effect becomes negative) implying that current account deficits increases the unemployment reinforcing the findings of Granger causality test. As mentioned in (Zachariadis and Pashourtidou, 2007, p.195), the existence of the cointegrating relationship between current account deficits and unemployment, we observe that shocks don not fade away and cerate permanent trace on the affected variables.

4-4. Variance decompositions

Based on the results of VECM Granger causality tests, we concluded that there is a unidirectional causality running from current account deficits and unemployment for the period of 2000Q1 and 2012Q1. As we mentioned in section 3.2 above, the VECM Granger

causality test only indicate the causality within the sample period, and does not allow us to capture the direction of causality among the variables beyond the sample period. According to (Abu-Bader and Abu-Qarn, 2008, p. 895), to capture the Granger causality among the variables beyond sample period, one has to portion the variance of the forecast error of a certain variable into proportions attributable to shocks in each variable in the system including its own, variance decomposition (VDC) allows us to detect the Granger causality beyond the sample period. Table 3 reports the results of VDC.

Table 3
Results of variance decomposition

Variance Decomposition of CAD: Period	S.E.	CAD	UEMP_SA
1	2397.657	100.0000	0.000000
2	3434.320	86.99121	13.00879
3	3957.173	87.49559	12.50441
4	4150.789	86.61750	13.38250
5	4572.782	88.45357	11.54643
6	4681.946	87.01459	12.98541
7	4703.782	86.26213	13.73787
8	4728.425	86.38283	13.61717
9	4864.061	82.58281	17.41719
10	5053.298	79.69543	20.30457
Variance Decomposition of UEMP_SA: Period	**S.E.**	**CAD**	**UEMP_SA**
1	0.005042	31.88404	68.11596
2	0.009639	59.05892	40.94108
3	0.013264	64.49566	35.50434
4	0.014937	67.01318	32.98682
5	0.015750	70.24765	29.75235
6	0.015772	70.30623	29.69377
7	0.015799	70.13394	29.86606
8	0.015881	69.56977	30.43023
9	0.016424	69.03509	30.96491
10	0.017230	69.12188	30.87812
Cholesky Ordering: CAD UEMP_SA			

The VDC results seem to validate the Granger test results. Current account deficits explain almost 70% of an innovation in unemployment while unemployment explains only 20% of an innovation in current account deficits in 10 periods. 79% of an innovation in current account deficits is explained by itself. Therefore, the results of the VDC are consistent with the Granger causality test and impulse response results.

5. Conclusion and Policy Discussion

In this paper, using quarterly data over 2000 to 2012 we investigated the (Granger-) causality relationship between unemployment and current account deficits in Turkey. Our results indicated that current account deficits explain a substantial fraction of the variation in unemployment and suggest the presence of strong unidirectional causality. We interpret these findings as evidence of the structural sources of unemployment being embedded under the deepening external fragility of the Turkish economy over the 2000s. The persistent unemployment rates of 9 – 10% over a decade –despite rapid growth, suggests that the problem is by no means conjectural, and rather structural.

Based on these observations we suggest a series of policy questions for further research.

- Exchange rate appreciation stands as a key obstacle discouraging employment friendly industrialization and patterns of growth. How to avoid currency appreciation in a time of capital mobility and excessive international credit in forms of "hot" finance?

- Related to this, over-emphasis on export-led growth with an over-reliance on foreign direct finance often lead to trap labour remunerations and decent job expectations to dual market structures with a minority enjoying rights of formal labour markets, at the expense of increased fragility and informality within a vast pool of unprotected/vulnerable workers. How to re-orient and balance the industrialization strategy across the warrants of external and domestic equilibrium?

- In many countries the gap between wage earnings and productivity of labour is widening, with a consequent fall of the labour share out of national income. Granted that much of this mechanism has to do with the desire to lower the unit costs of labour, social protection mechanisms and efficient waging institutions ought to be found to enable workers to capture their fair share of the fruits of their labor.

Finally,

- The issue of inflation targeting in the context of a de-regulated, open financial account within the free floating exchange rate regimes tend to create an environment of *inflation phobia* with a bias against the real sectors. Ironically, employment creation has dropped off the direct agenda of most central banks just as the problems of global unemployment, underemployment and poverty are taking center stage as critical world issues. The key problem is that the ongoing "financial globalization" appears primarily to redistribute shrinking investment funds and limited jobs across countries, rather than to accelerate capital accumulation across global scale (Akyuz, 2006; Adelman and Yeldan, 2000). Simply put, the world economy is growing too slowly to generate sufficient jobs and it is allocating a smaller proportion of its income to fixed capital formation.

- It is clear that the problem of unfriendly employment patterns of *speculative growth* ought to be tackled first and foremost by proper *macro* policies at large, rather than the much fashionable proposals of *micro-managerial reforms*.

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Appendix 1.

A-1-1. Unit Root Tests

The ADF tests based on the three following equations:

$$\Delta y_t = \alpha + \beta t + \delta y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \quad (1)$$

$$\Delta y_t = \alpha + \delta y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \quad (2)$$

$$\Delta y_t = \delta y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \quad (3)$$

In the ADF tests, null and alternative hypothesis are expressed as

$H_0 : \delta = o$

$H_1 : \delta \text{ p } o$

Here the null hypothesis (H_0) is the presence of a unit root and the alternative hypothesis (H_1) is the stationary of the series. The null hypothesis of a unit root is not rejected when the calculated ADF statistics for the variable exceed the critical value at the conventional significance level such as 5%. The distribution of the ADF statistic is non-standard and critical values are tabulated by MacKinnon (1991) are used.

When interpreting the results of the ADF test we have to aware of two crucial facts. Firstly, this test is very sensitive to incorrect establishment of the lag structure. Secondly, this test is often known as significant under-rejection (H. Gurgul and Lach, 2011). It is now well known that ADF test has a low power in rejecting the null of a unit root (Liang and Teng, 2006, p. 403). Therefore, in order to confirm the outcomes of the ADF test, PP and KPSS tests are also additionally carried out.

The PP test suggests the use of a specific nonparametric method for controlling for serial correlation when checking for unit root. The null hypothesis in this test once again refers to non-stationarity ($H_0 : \delta = o$) and alternative is $H_1 : \delta \text{ p } o$. In this test following equation is estimated:

$$y_t = \alpha + \delta y_{t-1} + \gamma(t - /T/2) + \varepsilon_t \quad (4)$$

where T is the number of observations and ε_t is error term which has zero mean but there is no requirement that it is serially uncorrelated or homogenous. Eq. (4) is estimated by OLS and t-statistics of the coefficient of lagged dependent variable y_{t-1} is corrected for serial correlation in error term by using the Newey-West procedure for adjusting the standard errors.

The PP test is used because it allows for milder assumptions on the distribution of error term (Kouakou, 2011, p. 3641). Also, as pointed out by B. Lin et al.(2012), although PP

tests uses the same models as ADF tests, it is considerably insensitive to the heteroskedasticity and the autocorrelation of the residuals of test equations. In other words, as pointed out in (Dritsakis, 2004, p. 255), in the case of and weakly autocorrelation and heteroskedastic regression residuals, PP is test is considered more robust. Also, it is believed that PP test is more powerful than the ADF test for the aggregate data.

However, since ADF tests and PP tests on small samples of data may be inefficacy, KPSS test is more effective for small samples when it chooses a lower lag truncation parameter. Also, since the null hypothesis under KPPS test is a trend stationary process and while the null hypothesis under the ADF and PP tests is the presence of a unit root, the use of KPSS test provides a good cross-check at conventional significance levels such as 5% (Le and Chang, 2013, p. 82). Because of these reasons, we implemented all unit root tests of ADF, PP and KPSS to assess the stationary time series.

According to Kwiatkowski et al. (1992), a time series consists of a deterministic trend, a random walk and a stationary error:

$$y_t = \delta t + r_t + \varepsilon_t$$

Where r_t is a random walk $r_t = r_{t-1} + \nu_t$. The ν_t : $iid(0, \sigma_\nu^2)$[2]. The null hypothesis state that y_t is stationary around a constant ($\delta = 0$) or trend stationary ($\delta \neq 0$). To carry out the KPSS test, first we should run a regression of y_t against a constant to test whether y_t is level-stationary or a constant plus a time trend for testing whether variable y_t is trend-stationary. To compute the sample value of LM test statistics after obtaining the residuals from the regression, one can use the following formula:

$$LM = T^{-2} \sum_{t=1}^{T} \frac{S_t^2}{S_{\varepsilon t}^2}$$

where $S_{\varepsilon t}^2$ is the estimated value of variance of ε_t and S_t is the sum of the residuals. The distribution of LM is non-standard and test is an upper tail test. One can find the limiting values from Kwiatkowski et al. (1992).

On the other hand, the existence of structural breaks in the series affects the outcome of unit root tests. ADF, PP, and KPSS tests tend to be biased in favour of the null of a unit root if there are structural breaks in the series (Tao and Green, 2012, p. 28). Zivot-Andrews (ZA hereafter, 1992) develop an alternative unit root test which can be used to evaluate the stationarity of series in the case of a structural break and that treat the occurrence of the break date as unknown. We therefore also performed the ZA test allowing for endogenous one-time break in intercept and/or trend. For this purpose, we use three models. These models are Model A which allows break in intercept, Model B which allows break in trend, and Model C which allows break in both intercept and trend. The null hypothesis in ZA test is that series contain a unit root with structural break or the coefficient of lagged dependent variable (α in Models below) is equal zero. The alternative is that series are trend-stationary with one-time break at an unknown time that is to be estimated. As indicated (Taha and Maslyuk, 2013), the ZA test allows estimating the following augmented regression equations:

[2] Stationarity implies that $\sigma_\nu^2 = 0$.

Model A: $y_t = \mu + \alpha y_{t-1} + \beta t + \theta DU_t + \sum_{j=1}^{k} \varphi_j \Delta y_{t-j} + \varepsilon_{1t}$ (5)

Model B: $y_t = \mu + \alpha y_{t-1} + \beta t + \gamma DT_t + \sum_{j=1}^{k} \varphi_j \Delta y_{t-j} + \varepsilon_{2t}$ (6)

Model C: $y_t = \mu + \alpha y_{t-1} + \beta t + \theta DU_t + \gamma DT + \sum_{j=1}^{k} \varphi_j \Delta y_{t-j} + \varepsilon_t$ (7)

where DU and DT are indicator variables for a mean shift occurring at possible break-date (TB) and for a shift in trend respectively and DU = 1 and DT = t-TB if t>TB and 0 otherwise; Δ is the first difference operator; ε_t is a white noise error term at time t; k is lag length. The selected break-date for each data series is TB where the t-statistics for the null hypothesis is minimized. Finally, the t-statistic for y_{t-1} can be used test the null hypothesis and the null hypothesis is rejected if the coefficient of y_{t-1} is significantly different from zero.

A-1-2. Johansen Co-integration Test

According to Engle and Granger (1987), if non-stationary time series have the same order of integration, for example order one and if these time series linear combination exist and stationary, which is integrated of order zero, then these time series are called cointegrated time series.

According to (Love and Chandra, 2005, p. 1161), once we found that the variables are non-stationary at their level and are stationary at their first differences, that is integrated of order one, we can find out whether they are cointegrated employing Johansen framework details of the method can be found in Johansen (1988) and Johansen and Juselius (1990). This method relies on the relationship between the rank of a matrix and its characteristic roots or eigenvalues (Gilmore and McManus, 2002, p. 81). Let x_t is be a vector of variables integrated of order one of dimension px1. The representation VAR of order k is given by

$$X_t = \mu + \Pi_1 X_{t-1} + \ldots + \Pi_k X_{t-k} + \varepsilon_t \quad (8)$$

where $\Pi_1 \ldots \Pi_k$ are (pxp) lag coefficients matrices, ε_t is a (px1) dimensional independently and identically distributed Gaussian error term with zero mean and non-singular variance-covariance matrix. μ is a vector of constants. Since X_t is non-stationary, the Eq. (8) can be rewritten in error-correction form

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_k \Delta X_{t-k+1} + \Pi X_{t-k} + \varepsilon_t \quad (9)$$

Where $\Gamma_i = -1 + \Pi_1 + \ldots + \Pi_i$ with i=1,…,k-1 and $\Pi = -(1 - \Pi_1 - \ldots - \Pi_k)$. The coefficient matrix Π contains information about the long-run relationships between variables in data vector ((Love and Chandra, 2005, p. 1162). From the Eq. (9), it is impossible to have relationship between a variable integrated order one, that is I(1) and a variable which is I(0).

As mentioned in (Kühl, 2010, p. 3), the Johansen approach adopts the idea of determining the rank (r) of the long-run coefficients matrix Π. Thus, by examining the rank of Π matrix, we can determine whether cointegration exits among X variables. There are three possible cases. First, the matrix has full rank, i.e. rank of Π equals p, all elements of X are stationary in levels and none of the series has a unit root. Second, if the rank of Π=0, i.e. Π is a null matrix meaning that all elements in the adjustment matrix have value zero, therefore, according to (Alam et al., 2012, p.219), none of the linear combinations are stationary and an unrestricted VAR model can be estimated to capture the short-run dynamics only. Third, if the rank of Π=r such that o<r<p, there are r cointegrating vectors or r stationary linear combinations. For example, if the rank of Π=1, there is a single cointegrating vector or one linear combination which is stationary such that the cointegrating rank matrix Π can be decomposed into matrices α and β so that $\Pi = \alpha\ \beta'$. The resulting matrix α contains the speed of adjustments and the matrix β contains the coefficients of cointegration relations. In this case X_t is I(1) but the combination $\beta' X_{t-1}$ is I(0).

Two likelihood ratio (LR) tests are used for detecting the presence of co-integrating vectors. The first is the trace test, which tests the null of at most r co-integrating vectors against the alternative that it is less than r. The LR statistics for trace test is the following equation:

$$\lambda_{trace} = -T \sum_{i=r+1}^{p} T \ln(1 - \hat{\lambda}_i)$$

where $\hat{\lambda}_i$ is the estimated values of the characteristic roots (eigenvalues) obtained from the estimated Π matrix and T is the number of usable observations after lag adjustments.

The second is the maximum eigenvalue test, which tests the null of r co-integrating vectors against the alternative of r+1. The LR statistics for maximum eigenvalue test is the following equation:

$$\lambda_{\max} = -T \ln(1 - \hat{\lambda}_{r+1})$$

Both test statistics are distributed asymptotically as χ^2 with p-r degrees of freedom. The maximum likelihood technique is used to estimate parameters of system. If variables in the system are not cointegrated, the rank of Π will be zero and so will be all the characteristic roots. Since ln(1) = 0, each of the expressions $(1 - \lambda_i)$ will equal zero in that case. According to (Abbas and Choudhury, 2012, p. 5), the trace test shows more robustness to skewness than maximum eigenvalue test.

Appendix 2.

A-2-1. Unit Root Test Results

Table A-1 reports the results of unit root tests based on ADF, PP and KPSS tests for the variables used in the analysis in levels and first differences. The optimal lag lengths of ADF test are chosen based Akaike information criterion (SIC) and optimal band-widths of PP and KPSS tests are determined based on Newey-West criterion.

Table A-1. Results of the unit root tests.

	Lag length	ADF	Bandwith	PP	Bandwith	KPSS	Conclusion
CAD	5	-1.35	9	-1.43	5	0.71**	I(1)
ΔCAD	4	-3.41**	19	-11.11*	47	0.50**	I(0)
UNEMP	1	-2.28	2	-2.15	5	0.45***	I(1)
$\Delta UNEMP$	0	-4.55*	1	-4.55*	3	0.26	

Notes: For all tests**,** a constant is included. For the ADF test, optimal lag lengths are determined by using AIC with a maximum lag of 10. For both PP and KPSS tests the spectral estimation method is the Bartlett kernel, while bandwidth is the Newey-West.
*Denotes significance at the 1% level.
**Denotes significance at the 5% level.
***Denotes significance at the 10% level.

In table A-1, for the current account deficits and unemployment series, the ADF and PP statistics cannot reject the null hypothesis of unit root at the all conventional significance levels while KPSS significantly reject the null hypothesis of stationary, implying that all series are not stationary in their levels. For the first differences of the current account deficits and unemployment series, the ADF and PP tests consistently indicate that they are stationary at the all conventional significance levels while KPSS test only fail to reject the null hypothesis of stationary at the 1% significance level. Based on the above unit root tests, we concluded that all variables are non-stationary in their levels but become stationary after taking the first difference. In other words, we conclude that all series are integrated of order one which is I(1) at the 1% level of significance.

Besides the unit root tests of ADF, PP and KPSS, to avoid the false identification of the order of integration, we also implemented the ZA tests for existence of unit roots in the variables in the study, allowing this time for the presence of one structural break. According to Dergiades et al. (2012, p. 6), not taking into account of for possible structural break may give rise to the so-called Perron phenomenon or the converse Perron phenomenon. The results of the three alternative specifications of the ZA test presented in table A-2.

Table A-2. Results of ZA unit root tests.

	Model A			Model B			Model C		
Variable	t-stat.	k	Break-date	t-stat.	k	Break-date	t-stat.	k	Break-date
CAD	-4.57	2	**2008Q3**	-4.07	2	**2009Q2**	-5.93	2	**2008Q4**
UNEMP	-2.61	1	**2010Q2**	-2.92	1	**2009Q4**	-4.03	1	**2008Q4**
Critical values at the 1%, 5%, and 10%	-5.34,-4.93, -4.58			-4.80,-4.42, -4.11			-5.57, -5.08, -4.84		

Clearly for all models (A, B, and C), the ZA tests fail to reject the null hypothesis of unit root with one structural break in levels, at all the conventional significance levels. Therefore, the ZA tests find no additional evidence against the ADF, PP, and KPSS unit root tests. For this reason, we concluded that the ZA tests results corroborate the findings of the other unit root tests that all variables in the study integrated of order one, i.e. I(1). Note that the same order of integration, i.e. one, is a pre-requisite when Johansen method is used for testing for cointegration and then causality. Thus, we can proceed Johansen cointegration test.

A-2-2. Johansen Co-integration Test Results,

Before carrying out the cointegration tests, we have to specify optimal lag length of the variables, since the cointegration tests results are sensitive to the choice of lag length. To determine the optimal lag length, we estimated VAR in levels and used both Akaike Information Criterion (AIC) and Final Prediction Error (FPE) criterion (see Table 3), since the information criteria do not favor the same lag length. We used both maximum eigenvalue ($\lambda_{\max}$) and the trace (λ_{trace}) test statistics to test for a cointegrating relationship among variables.

Table A-3. Results of lag length selection from unrestricted VAR model.

Lag	LogL	LR	FPE	AIC	SC
0	-278.0653	NA	5922.956	14.36232	14.44763
1	-224.8388	98.26432	474.7277	11.83789	12.09382*
2	-219.2498	9.744881	438.5545	11.75640	12.18295
3	-217.7839	2.405669	501.9295	11.88635	12.48353
4	-209.0750	13.39820	397.8398	11.64487	12.41267
5	-199.6991	13.46283	306.3916	11.36919	12.30761
6	-197.4320	3.022842	342.1531	11.45805	12.56709
7	-193.3356	5.041731	350.9880	11.45311	12.73277
8	-187.8992	6.133421	339.9227	11.37944	12.82973
9	-186.3422	1.596837	407.3527	11.50473	13.12564
10	-172.2327	13.02423*	261.0300*	10.98629*	12.77782

* indicates lag order selected by the criterion

The results of the Johansen maximum likelihood cointegration tests are presented in Table A-4.

Table A-4 The results of Johansen co-integration test.

Trace				**Max Eigenvalue**			
H_0	H_1	Statistics	5% critical values	H_0	H_1	Statistics	5% critical values
r=0	$r \geq 1$	28.83017	15.49471	r=0	$r = 1$	27.06977	14.26460
$r \leq 1$	$r \geq 2$	1.760396	3.841466	$r \leq 1$	r=2	1.760396	3.841466

The Johansen co-integration tests results in Table 4 show that the Johansen test identifies one co-integrating vector between CAD and UNEMP series under the both the Trace statistics and Max Eigenvalue statistics at 5% significance level. Therefore, CAD and UNEMP are co-integrated and VECM is the appropriate specification for the full-sample Granger causality tests.