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# The Greek Hyperinflation Revisited

By

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## Abstract

The objective of this paper is to gain an insight into the Greek hyperinflation that occurred during the period 1941-1946. In doing so, a relatively novel data-set in conjunction with the bound testing approach to cointegration and error correction models developed within the autoregressive distributed lag (ARDL) framework, shed additional light on the underlying long-run relationship between money supply and inflation. Granger causality tests between money supply and prices are also conducted in the effort to ascertain the direction of causality between money supply and the (hyper)inflation rate.

**Key Words:** *Hyperinflation, Cointegration, ARDL*

**JEL classification numbers:** *C32, E31, N10, O11*

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## 1. Introduction

The Greek hyperinflation became widely known from the publication of Philip Cagan's (1956) seminal paper which ranked the Greek hyperinflation second from a list of seven hyperinflations with first being that of Hungary during WWII. According to Cagan's (1956, p. 21) working definition, hyperinflation starts in the month when the inflation rate exceeds the 50% threshold and remains at this level; the phenomenon ends from the last month after which inflation remains below the 50% borderline for at least twelve months. This definition of hyperinflation was derived from the demand for money for transactions purposes and despite of its empirical character it has prevailed not only in the empirical but also in the theoretical literature.

We know that hyperinflations occur during abnormal periods such as those that characterize countries under occupation or during civil wars and, in general, in countries with weak governments and lack of social cohesion. Under such circumstances, governments lose fiscal discipline and they use their seignorage as a means to finance their expenditures. As a result, hyperinflations occurred in countries such as the former Yugoslavia during 1992-1994; former Soviet republics, such as Ukraine, during 1993-1995, Russia during 1992-1994, among others; Latin American countries where governments were financing their expenditures merely by printing money and, of course, African countries that experience civil wars, Mozambique is the most recent example. As for the Greek hyperinflation, we argue that Cagan's (1956) data set was neither representative of the severity of the phenomenon nor of its longevity. In a similar fashion, studies carried out by Sargent and Wallace (1973), Makinen (1986 and 1988) and Karatzas (1986) generated evidence that appear to be suffering from the same weakness. The study by Anderson, Bomberger and Makinen (1988) is probably the most comprehensive one touching on the reliability of the data series that have been used in Cagan's original study; nevertheless Anderson *et al.* (1988) do not make the necessary adjustments so as to describe the phenomenon of hyperinflation to its full extent and also to address the question of the stability of the demand for money. Furthermore, the econometric techniques that have been used by Cagan (1956) or by Sargent and Wallace (1973) and more recently by Anderson *et al.* (1988) are considered outdated, with today's standard. In addition, these techniques suffer from a number of limitations that may give rise to spurious results.

In this paper not only do we present a more accurate description of the phenomenon of the Greek hyperinflation, but we also employ a recently advanced

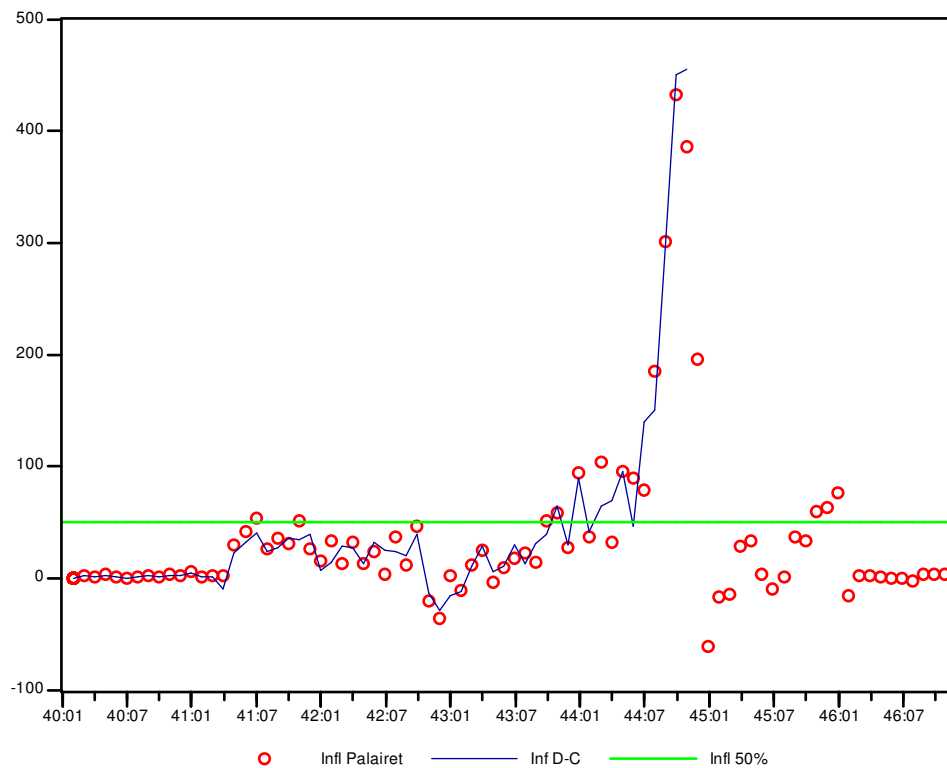
econometric technique i.e. the Autoregressive Distributed Lag (ARDL) approach, in an attempt to gain a more comprehensive insight into the analysis of hyperinflation. The rest of the paper is organized as follows. Section 2 touches on the historical background and re-evaluation of the existing evidence. Section 3 describes the ARDL model of hyperinflation in the case of the Greek economy, while section 4 presents and discusses the results of the analysis. Finally, section 5 provides a summary coupled with some concluding remarks.

## **2. Historical Background**

The Greek hyperinflation emerged into prominence mainly due to the seminal paper by Cagan (1956) within which countries experiencing unstable and accelerated levels of inflation were investigated. The data used in his study come from Delivanis and Cleveland's (1949) book (henceforth D-C) on the Greek economy mainly during the years of occupation (April 1941- October 1944). A major drawback of D-C's data-set is that the price data might have undermined both the seriousness of the phenomenon of hyperinflation as well as its time span. In view of the latter, an alternative data set compiled by Agapitides (1945) and extended by Palairat (2000) can be used to gauge the extent of the underlying bias permeating the D-C data-set. In particular, the D-C price index includes goods and services of the Athens area alone, whereas Agapitides reports a price index which comprises a basket of goods corresponding to a lower calorie diet and the coverage is broader than that of D-C for it includes Athens and Piraeus areas. As a result, Agapitides's price index is much more representative and therefore reliable than the price index reported by D-C. Furthermore, Agapitides reports detailed data on services, and from the housing services the rent component is reported separately. This is extremely important, because rent occupies a relatively large share in the cost of living and this may distort the true cost of living of the people during the time of our investigation, because of the rent control that was imposed already in the 1940 and continued during the years of occupation and even the years after. As a matter of fact, the price index without the rent is 15% higher than the index with the rent, the difference increases up to 47% and remains at this level until the end of 1944 and continues to be 15% higher for the rest of the period. Naturally, the price index without the rent and the associated with it inflation rate will be higher than the price index and inflation rate which include the rent.

Figure 1 below depicts the inflation rate used in Palairret (2000) for the most turbulent period of the Greek monetary history, *i.e.* from 1940:1 until 1946:12. In the same figure we plot the 50% borderline inflation rate suggested by Cagan (1956), whereas the D-C inflation rate spanning the period 1940:1-1944:11 that was used by Cagan (1956) is also illustrated in the same figure. Evidently, the hyperinflation in Cagan's study spans for a shorter period starting in November of 1943 and ending 12 months later. In fact, D-C data end in November 10<sup>th</sup> of 1944.

Figure 1. Inflation Rate 1940:1-1946:12



Given that the inflation rate in studies of the demand for money normally is a measure of the opportunity cost of holding money, it follows that what is sought is the rate of increase in the price or prices of the major alternatives to money as a store of value. Lack of knowledge concerning other possible alternatives has led investigators to assume that the price of non-perishable commodities are the major alternatives to money in the majority of hyperinflations. Finding a suitable index however to measure the rate at which they rise in price has proven a difficult task. An index either of wholesale or consumer prices has been frequently used as a proxy, even though it

may contain certain services such as housing and utilities, which cannot be stored and these prices may rise much more slowly than the prices of nonperishable commodities. Luckily, Agapitides' (1945) price index is detailed and reports separately the rent component of the price index. Palairret (2000) running a regression between the real money supply against the inflation rate found that this borderline inflation rate is 42 percent, which is somewhat lower than the Cagan's 50 percent borderline in the case of the Greek data.

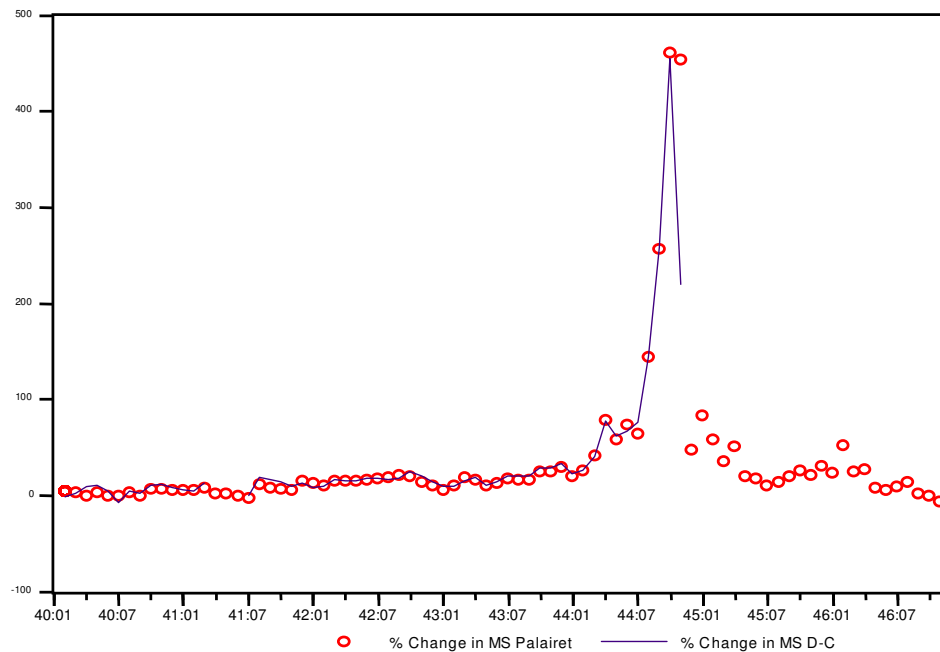
As for the supply of money, we find that the D-C data set refers to the money supply which is restricted to the number of notes issued by the Bank of Greece (the central bank) at the end of the month. Anderson *et al.* (1988) however have argued that a somewhat more reliable data series is reported by Agapitides (1945) where one can find monthly data for the number of notes issued by the Bank of Greece which were pumped into circulation. Cagan's (1956) study based on D-C data set finds a time period of hyperinflation shorter than that derived in Agapitides (1945). Furthermore, the concept of money supply used in the hitherto studies is too narrow for it is restricted to the number of notes issued by the Bank of Greece, a bias which is serious in the period during which hyperinflation reaches its crescendo. Palairret's (2000) data on the number of notes issued together with the sight deposits reported by the same source coupled with some insight information from internal sources of the Bank of Greece give us a more reliable estimation of the supply of money during the period under investigation.<sup>1</sup>

Moreover, the money supply in its narrow sense (M0) *i.e.* the number of notes in circulation—the proxy for the money supply in Cagan (1956) and Anderson *et al.* (1988)—has come in for a lot of criticism in that the data come from the same source *i.e.*, the Bank of Greece and, therefore, the two series (*i.e.*, D-C and Agapitides, 1945) are identical up until August of 1942. From September of 1942 until October of 1944, the money stock in D-C exceeds the notes in circulation reported by Agapitides; in particular, for the time period of September 1942 until December 1942 the D-C data on issue of notes is 11% higher than that of Agapitides, whereas the difference drops by one percentage point during the twelve months of 1943. From January 1944 until July 1944 the D-C index is higher by 26% and from August through October 1944 the difference of the D-C increases by 900%. According to Anderson *et al.* (1988) Agapitides's figures are much better than those of D-C for two reasons:

- (a) The practice of the Bank of Greece was to report “notes in circulation” all notes issued including those that were kept in the vaults of its various branches (non-circulating notes). It seems very likely that the Bank of Greece in anticipation of future inflation had the necessary notes ready to pump them into circulation, if such a need arose. As a consequence, the issued notes would exceed those actually in operation and this difference is very likely to be depicted in Agapitides’s data.
- (b) We know that in November 1944 we had a reform of the monetary system of Greece and the data that D-C report money supply of 121 million new drachmas. However, we also know that in fact 222 million drachmas had been converted. This fact indicates that after November 11, 1944 large amounts of old drachmas were issued something that is consistent with the money supply data reported by Agapitides.

Palairret (2000) offers a much simpler and to our view convincing explanation. More specifically he notes that the source of differences of the two data series may stem from the fact that D-C refer to the end of month data and Agapitides’s data refer to the 15<sup>th</sup> of each month. Palairret (2000) having to choose between the two alternative estimates opted for the end of month data combining other sources, whenever D-C did not provide data as for example in the months of April and May 1941 and for October 1944 and after November 10<sup>th</sup> of 1944 (for details Palairret, 2000, p. 113). Furthermore, Palairret (2000) provides data of the sight deposits of the public in private banks, which when added to the notes in circulation give the measurement of money supply, M1. As a consequence, we have a much more reliable index of money supply spanning a period of time longer than that of D-C. In fact, Palairret (2000) essentially adds the sight deposits to the money supply estimates of D-C and extended the data up until 1946:12. Figure 2 below, displays the percentage change of the two definitions of money supply. Even though, at first glance, the two series appear to be fluctuating in a rather similar fashion, and hence no substantial effects are envisaged at least from an econometrics point of view, there is an advantage in Palairret’s data set in a sense that it includes a broader definition of money supply as well as a much longer time span, allowing thus a more reliable discussion of the longest hyperinflation ever occurred at least in Europe.

Figure 2. Percentage Change in Money Supply



### 3. Empirical Investigation

#### *The model*

Despite the fact that the phenomenon of hyperinflation has been discussed extensively in the literature nevertheless there is no particular theory to explain it adequately. The general view is that what causes inflation causes hyperinflation. In looking into the underlying relationship between the logarithm of the price index ( $\pi$ ) and the logarithm of the money supply ( $m$ ) a generic long-run model is fostered and effectively applied in the following form:

$$\pi_t = \beta_0 + \beta_1 m_t + \varepsilon_t \quad (1)$$

Where,  $\beta_0$  is the constant;  $\beta_1$  is the slope coefficient;  $\varepsilon$  is the error term and  $t$  is time.

For the econometric analysis monthly time series data has been collated for Greece spanning from 1941:01-1946:1 (Palairet, 2000). It is worth noting at the outset that this econometric methodology is appropriate when there is a rather a larger number of observations spanning a long time period. However, for episodes of hyperinflations as Hakkio and Rush (1991) have argued the “long run” may in some cases “be a matter of months”. In fact, for hyperinflation episodes, a few months are all that matters.



### Methodology

The present study by employing cointegration techniques and error correction modelling (ECM) attempts to unravel the ‘mystery’ of hyperinflation. Cointegration analysis provides potential information on long term equilibrium relationship between inflation and money supply. The ECM on the other hand, as a tool of analysis overcomes the problems of spurious regression through the use of appropriate differenced stationary residuals in order to determine the short term adjustments in the model. Given the fact that most time series generally exhibit a non-stationary pattern in their levels, unit root testing will be carried out in order to determine the degree of stationarity. The econometric methodology consists of the following steps: Firstly, we check the series to determine the order of integration<sup>2</sup>. The Augmented Dickey-Fuller (1979) test has been extensively used in empirical studies when determining the order of integration<sup>3</sup>. More recent studies however generated evidence indicating that in the presence of a structural break, the standard ADF tests are biased towards the non rejection of the null hypothesis (Perron, 1989). In view of the latter a number of scholars tried to overcome the problem by proposing a very specific treatment *i.e.* the endogenous determination of the break using the existing data<sup>4</sup> (see for instance, Banerjee, Lumsdaine and Stock, 1992; Zivot and Andrews, 1992; Perron and Vogelsang, 1992; Perron, 1997 and Lumsdaine and Papell, 1997)<sup>5</sup>.

Given the existing criticism surrounding the conventional ADF method we proceed to testing whether the unit root tests for the variables were biased because possible breaks in the series were ignored (Perron 1997, Zivot and Andrews (1992)). Drawing on Perron’s structural break test the estimating regression assumes the following general form:

$$y_t = c + aDU_t + \beta t + \gamma DT_t^* + \zeta D(T_B)_t + \mu y_{t-1} + \sum_{i=1}^p \phi_i \Delta y_{t-i} + \varepsilon_t \quad (2)$$

where  $DU_t$  is the intercept dummy,  $DU_t = 1$  if  $(t > T_B)$  and 0 otherwise; where  $DT_t$  is a slope dummy representing a change in the slope of the trend function;  $DT_t = t - T_B$  (or  $DT_t^* = t$  if  $t > T_B$ ) and 0 otherwise; where  $(DT_B)$  is the one time break dummy,  $(DT_B) = 1$  if  $t = T_B + 1$ ; and  $T_B$  is the break date. Prior to determining the date of a structural break endogenously, it is imperative that the date which minimizes the Dickey Fuller t-statistic for testing the null hypothesis of a unit root ( $a=1$ ) is selected. Subsequently, the date of the structural break is chosen such that the value of  $|t_\gamma|$  is

maximized<sup>6</sup>. Finally, through utilizing a general to specific approach the value of the lag truncation parameter  $k$  will be determined<sup>7</sup>. In passing, it should be noted that the model developed by Perron (1997) is slightly different from the one coined by Zivot and Andrews (1992) in that the latter is devoid of the one time break dummy ( $DT_B$ ).

Having established the order of integration we secondly engage in testing the cointegration of the series utilizing the bounds testing approach within the ARDL framework. In recent years, reams of academic papers have been produced proposing different methodologies on how to investigate long-run equilibrium between time-series variables. On the univariate front, cointegration techniques such as the ones by Engle and Granger (1987) and Phillips and Hansen's (1990) have been applied. As for multivariate cointegration, Johansen (1988) and Johansen and Juselius (1990) full information maximum likelihood procedures are extensively used in empirical studies. A relatively new procedure, the Autoregressive Distributed Lag (ARDL), introduced originally by Pesaran and Shin (1999) and further extended by Pesaran *et al.* (2001) and Narayan (2005) also deals with single cointegration. This method is thought to have certain econometric advantages over other single cointegration procedures. More specifically, endogeneity problems and inability to test hypotheses on the estimated coefficients in the long-run associated with the Engle-Granger method are avoided; the long and short-run parameters of the model are estimated simultaneously; all variables are assumed to be endogenous; it also obviates the need to establish the order of integration amongst the variables, *i.e.*, the Pesaran *et al.* (2001) method could be implemented regardless of whether the underlying variables are  $I(0)$ ,  $I(1)$ , or fractionally integrated.<sup>8</sup>

To implement the ARDL approach, equation (1) is transformed to a conditional error correction version of the price level and its determinants:

$$\Delta\pi_t = a_0 + \sum_{i=1}^p \beta_{1i} \Delta\pi_{t-i} + \sum_{i=1}^p \beta_{2i} \Delta m_{t-i} + \beta_3 \pi_{t-1} + \beta_4 m_{t-1} + \varepsilon_t$$

(3)

The first part of equation (3) with  $\beta_1$ , and  $\beta_2$  representing the short run dynamics of the model, whereas the second part with  $\beta_3$ , and  $\beta_4$  represents the long run relationship,  $\Delta$  is the first difference operator and  $p$  is the optimal lag length.

Next, the joint hypothesis that the long-run multipliers of the lagged level variables are all equal to zero against the alternative that at least one is non-zero will

be tested. If a cointegrating relationship exists then the null hypothesis should be rejected. The long run relationship amongst the variables is tested by means of bounds testing procedure coined by Pesaran *et al.* (2001). This procedure is based on the *F-test* or Wald-statistics and is the first stage of the ARDL cointegration method. A joint significance test that implies no cointegration is also performed. The *F-test* used for this procedure by performing the Wald test has a non-standard distribution, whose asymptotic critical values are provided by Pesaran *et al.* (2001). Further research on this area however has produced evidence on the basis of which the critical values are inappropriate whenever the sample size is small, or in other words when annual macroeconomic variables are involved (Narayan, 2005).

A number of regressions have been estimated in an attempt to obtain the optimal lag length for each variable. Once a long-run relationship is established, the long-run estimates can be obtained using the following ARDL specification:

$$\pi_t = \beta_0 + \sum_{i=1}^p \beta_{1i} \pi_{t-i} + \sum_{i=0}^q \beta_{2i} m_{t-i} + u_t \quad (4)$$

The order of lags in the ARDL model are selected by either the Akaike (AIC) selection criterion or the Schwartz Bayesian Criterion (SBC) before the selected model is estimated by ordinary least squares. For monthly data we chose 3 lags. From this, the lag length that minimizes SBC is selected.

Finally, the speed of adjustment to equilibrium level after a shock is captured by the error correction representation which is conveyed in the following form:

$$\Delta \pi_t = \beta_0 + \sum_{i=1}^p \beta_{1i} \Delta \pi_{t-i} + \sum_{i=1}^p \beta_{2i} \Delta m_{t-i} + \lambda EC_{t-1} + \varepsilon_t \quad (5)$$

where  $\lambda$  is the speed of adjustment; *EC* is the error correction component<sup>9</sup>, defined as:

$$EC = \pi_t - \beta_0 - \sum_{i=1}^p \beta_{1i} \pi_{t-i} - \sum_{i=0}^q \beta_{2i} m_{t-i} \quad (6)$$

Finally, given the order of integration of the underlying variables, an exploration of the causal dimension through Granger Causality tests will provide an indication as to the nature of causality between the two variables.

#### 4. Empirical findings

##### *Unit roots*

The initial step in analyzing the time series data properties, is to test for unit roots by applying the Augmented Dickey-Fuller (ADF) as well as Perron's (1997) and Zivot and Andrews' (1992) structural break tests. A quick inspection of the ADF results displayed in Table 1 below suggest that we can treat the underlying time series as I(1) variables<sup>10</sup>. Moreover on the basis of Perron's and Zivot and Andrews structural break test obtained the null hypothesis of a unit root is rejected at the 5 and 10 percent level of significance.

*Table 1: Unit Root Tests*

variables	ADF		Perron (1997)			Zivot and Andrews (1992)		
	levels	First Dif.	T <sub>B</sub>	k	t <sub>r</sub>	T <sub>B</sub>	k	t <sub>r</sub>
<i>II</i>	-0.765	-3.401*	1944.9	2	-5.322*	1944.9	2	-5.218*
<i>M</i>	-0.819	-2.949*	1944.9	1	-4.693**	1944.9	1	-4.982**

Notes:(\*), (\*\*) denote significant 5% and 10% tests respectively ; T<sub>B</sub> denotes the break date implied by t<sub>γ</sub> and t<sub>α</sub>. The critical values for the 1, 5 and 10 percent significant levels of the t statistics are -5.57, -5.08 and -4.82 for Zivot and Andrews' test and -5.57, -4.91 and -4.59 for Perron's test respectively.

##### *Cointegration tests*

On the basis of the bounds framework to cointegration the *F-statistics* should be compared to the critical values generated for specific sample sizes. Each variable in equation 2 is taken as a dependent variable in the calculation of the *F-statistics*. In particular, when inflation is the dependent variable the value of the *F-statistic* obtained is  $F_{II}(\pi \setminus m) = 7.765$  which is higher than the upper bound critical value of 4.363 at the 5% level of significance. The latter suggests that the null hypothesis of no cointegration can not be accepted. In contrast, the story is different when the dependent variable is money supply. More specifically, the *F-statistic* is found to be  $F_M(m \setminus \pi) = 3.542$ , which is lower than the upper bound critical value at the 5% level.

*Table 2: Bounds test for cointegration*

<i>T</i>	95% level	<i>Calculated F-statistic</i>	
	<i>I</i> (1)	$F_{II}(\pi \setminus m)$	$F_M(m \setminus \pi)$
57	4.363*	7.765	3.542

Notes: (\*) critical value bounds of the *F-statistic*: intercept and no trend. Critical value is taken from Narayan (2005), p. 1987.

Since inflation and money supply are cointegrated the long-run model using the ARDL specification (*i.e.*, equation 3) was estimated. In an attempt to find the optimal length of the level variables of the long-run coefficients, lag selection criteria based on AIC, and SBC were employed such as  $i_{max}=3$ . The yielding evidence suggests that there is a strong correlation between money supply and inflation over the sample period. In addition, the short term elasticities are found statistically significant at the 5% level reflecting thereby the existing relationship between the scrutinized variables.

*Table 3. ARDL Estimation results*

<i>Long-run elasticities (dependent variable is <math>\pi_t</math>)</i>		
<i>Constant</i>	$m_t$	
7.2(0.635)	0.067*(7.67)	
<i>Short-run elasticities (dependent variable is <math>\Delta\pi_t</math>)</i>		
<i>Constant</i>	$\Delta m_t$	$EC_{t-1}$
0.014(0.823)	0.64*(2.56)	-0.59*(-3.623)

Notes: t-ratios are given in parenthesis; (\*) significance at 5% level.

As for the coefficient of  $EC_{t-1}$  this is found to be statistically significant confirming the existing long run relationship between the variables. More specifically, the negative and strongly significant error correction component indicates a relatively speedy adjustment *i.e.* about 59% of the disequilibria of the previous month's shock adjust back to the long run equilibrium in the current month.

*Table 4. Pairwise Granger Causality Tests*

Null Hypothesis:	F-Statistic	Probability
$\Delta\pi$ does not Granger Cause $\Delta m$	5.30087	0.00031
$\Delta m$ does not Granger Cause $\Delta\pi$	3.09152	0.00970

In addition, Pesaran and Pesaran (1997) argued that it is extremely important to ascertain the constancy of the long-run multipliers by testing the above error-correction model for the stability of its parameters. The commonly used tests for this purpose are the cumulative sum (CUSUM) and the cumulative sum of squares (CUSUMQ), both of which have been introduced by Brown *et al.*, (1975). Figures 3 and 4 below, display the results of CUSUM and CUSUMQ tests, respectively. In both figures the dotted lines represent the critical upper and lower bounds at the 0.05 level of significance. The visual inspection of Figures 3 and 4, reveals that there is no evidence of parameter instability as the CUSUM and the CUSUMQ lie within the upper and the lower bounds.

Figure 3. CUSUM CUSUM of Squares tests

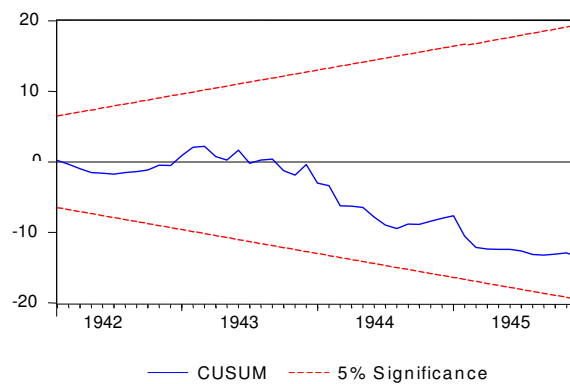
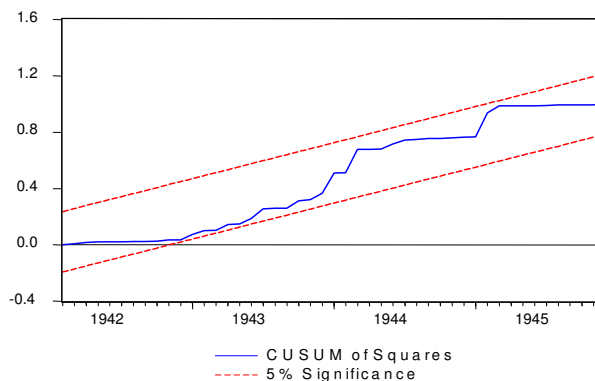


Figure 4. CUSUM of Squares test



Finally, it is worth noting that the Granger causality tests give another dimension to the existing relationship implying a bi-directional feedback. A result

which differs from that of Sargent and Wallace (1973), who found one way causality running from the growth of money supply to inflation, and the Greek hyperinflation (along with Hungarian II) was an exception to a host of other countries that were reported in Cagan's study. By contrast, the study by Anderson *et al.* (1988) by correcting some data found that the one way causality was rather from the inflation to the growth in money supply and so the explanation of the Greek hyperinflation brought into line with the other hyperinflations.

### **5. Concluding remarks.**

This paper has sought to re-examine the extent to which there is a long-run equilibrium relationship between the money supply and the price level in the Greek economy during a period permeated with hyperinflation. In order to minimize the bias associated with the small sample size and the associated time period, the ARDL approach to cointegration has been fostered as the latter is thought to be more efficient than the standard cointegration techniques. Furthermore, it is worth noting that the long-run is a rather relative concept and for the people experiencing occupation and hyperinflation even a period of a few months was considered long enough. The presence of cointegration has the intuitive meaning that although prices and money may both increase dramatically, nevertheless they tend to move together. The econometric evidence obtained lends support to the existence of cointegration between money supply and the price level indicating thus the existence of a rather stable money demand function during the period of hyperinflation. On the other hand, the error correction term suggests that whenever the two variables are out of equilibrium the equilibrium is restored in a relatively speedy way. In addition, the bi-directional causality suggests that even in exceptionally hyperinflationary periods the money supply may be an endogenous variable. This result is of no surprise since the money supply in conditions of hyperinflation plays an instrumental role in financing fiscal deficits. The latter does not necessarily imply that the money supply does not affect the price level, but rather that the relation between the two variables although stable is nevertheless too complex than is usually thought.

To sum up, our econometric analysis based on the ARDL approach confirmed the underlying relationship between the money supply and the price. At the same time however, it should also be noted that the Greek hyperinflation cannot be simply attributed to the unrestrained increase in money supply. Other factors of hardly

monetary nature should also be seriously taken into account. Finally, given the generic nature of the analytic framework utilized one can assume that invoking the latter can be proven useful enough in explaining not solely old but also recent episodes of hyperinflation in East Europe, Latin American and African countries. Gaining therefore an insight into hyperinflation and its dynamics is the ultimate challenge for all economic agents in an ever so fickle global economic environment.



## NOTES

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<sup>1</sup> Cagan (1956) points out the desirability of having access to a somewhat broader definition of money supply, such as M1 that he used in his study of the other hyperinflations. In particular, he notes: “Nevertheless, little stock can be placed in figures of such limited coverage. Furthermore, data on deposits and changes in real income are apparently nonexistent. Bank deposits should not be dismissed as entirely insignificant, though their effects in the other hyperinflations were minor, because deposits in Greece were as large in value as the quantity of bank notes in circulation during the hyperinflation” (Cagan, 1956, p. 106).

<sup>2</sup> It has been argued that time series data in a seasonally unadjusted form are preferred to their seasonally adjusted counterparts, since the filters used to adjust for seasonal patterns often distort the underlying properties of the data (see Davidson and MacKinnon 1993, for further evidence). It is also worth noting that macroeconomic time series can typically be described as  $I(1)$  with a deterministic seasonal pattern superimposed (Osborn, 1990). More specifically, Osborn (1990) found that only five out of thirty UK macroeconomic series required seasonal-differencing to induce stationarity. In the undertaken study we felt that getting bogged down into technical issues pertaining to seasonal unit root testing would not serve the purpose and rationale of the paper.

<sup>3</sup> The findings of Nelson and Plosser (1982) spawned a new wave of research to emerge on the unit root hypothesis. More specifically, the traditional notion that the current shocks only have a temporary effect and the long-run movement in the series is unaltered by such shocks was challenged by Nelson and Plosser (1982) who sustained that random shocks are bound to have permanent as distinct from transitory effects on the long-run level of macroeconomics.

<sup>4</sup> It should be stressed that the proposed endogenous tests came in for a lot of criticism for their treatment of breaks under the null hypothesis (see for instance, Lee and Strazicich, 2003).

<sup>5</sup> The advantages emanating from utilizing the procedure for testing the unit root hypothesis, which allows for the possible presence of the structural break, are twofold in a sense that it

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generates results free from biasness towards non-rejection as well as it traces the possible presence of structural break.

<sup>6</sup> We will select the breakpoint using the maximum of the absolute value of  $t_\gamma$ .

<sup>7</sup> As the scope of the paper is far from getting bogged down into the technicalities of the existing stationarity approaches the interested reader can gain a more insightful account of the processes adopted in Perron (1997) and Zivot and Andrews (1992).

<sup>8</sup> One major drawback of the ARDL approach to cointegration is that it fails to provide robust results when dealing with  $I(2)$  variables.

<sup>9</sup> It should be stressed that the standard diagnostics and stability tests have also been performed to test the validity of our model but not given for reasons of clarity and economy in space.

<sup>10</sup> It should be stressed that for the ADF tests, the lag length is based on the SBC.

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