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Long-Run Monetary Neutrality In Jordan

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Abstract

This study tests the hypothesis of long-run neutrality of money in Jordan for the period from January 1990 through August 2002. Money supply measures, namely M1 and M2, are used as monetary indicators, and the index of industrial production (1997 = 100) is used as a measure for the economic activity. The formal test of independence suggested by Koch and Yang provides no evidence against the neutrality of money hypothesis in the short-run as well as in the long-run. The empirical results are robust to three lag lengths (24, 30, and 36 months), and two different money measures (M1 and M2).

1. Introduction

Despite the large amount of empirical work on the impact of monetary policy on output, monetarists assert that government injections of money into an economy have a certain neutral effect in the long run. This aphorism implies that changes in money stock eventually change nominal prices and nominal wages, leaving important real variables; like real output, real consumption expenditures, real wages, and real interest rates unaffected. Since decision making process is based on real factors, the long-run effect of injecting money into the economy is often described as neutral. How long such a process takes, and what might happen in the meantime, are hotly debated questions (Haug and Lucas 1997, Bullard 1999, Leong, McAleer and Maki 1997 and 2000).

By mid of 1970's, the Jordanian economy began to suffer from serious imbalances. The rapid imbalance between aggregate domestic demand and aggregate supply was reflected in a worsening of its external payments' position and a rise in relative prices. The gap between aggregate demand and aggregate supply was bridged by imports, which were financed by the dramatic increase in workers' remittances and the unrequited transfers in the post of 1973 war. This beneficial supply shock did not last long. The sudden unfavourable decline in the oil price recorded in the beginning of 1980's adversely affected the economies of all countries in the region.

The increasing distortions in a declining growth rate of real output and a heavy external debt in Jordan have attracted the attention of international aid agencies and policy circles alike (Alkhatib 1995a). At the end of 1988, the exchange rate of the Jordanian dinar against the U.S. dollar dropped by about 50%, and the outstanding external debt exceeded US\$ 7 billions. More ominously, there were indications that Jordanian external debt has surpassed the capacity of the Jordanian economy.

At the beginning of 1990's, serious attempts to correct the economic imbalances have been suggested by the IMF and the World Bank. The broad objectives of Economic Adjustment Programs (EAPs) were the attainment of a viable balance of payments, a sustainable economic growth, and a low inflation rate. The formulated policies to achieve these targets included tight monetary policy, and some sectoral reforms aimed at liberalizing markets and remedying structural imbalances in various economic sectors (Khan 1990).

A question that is frequently raised regarding the EAPs is whether such programs are effective in achieving their broad objectives. What do people know about the effect of EPAs? How does output respond to policy measures? Some writers argue that these programs do little in the way of improving the economic situation. Others assert that these programs worsen the economic situation by causing stagflation (Summers 1993).

The purpose of this paper is to examine the dynamic response patterns of output to money movements using Jordanian data for the period starts from January 1990 through August 2002. During this period, the Central Bank of Jordan intensified its indirect intervention in the money market by issuing certificates of deposits as a saving instrument with the objective of regulating domestic liquidity and strengthening the position of the Jordanian dinar. During the period 1991-1998, interest rate was kept high and above 12.5%. As a result of a package of actions implemented by the Central Bank, money supply narrowly defined (M1) grew by about 4.2% on average..

There are at least two reasons to investigating the dynamic relationship between money growth rate and the growth rate of output during this period. First, it provides relevant information about the direction of causality between these two series, because most of the efficient estimation techniques are invalid unless causality is unidirectional. Second, the empirical testability of these issues is important for policy formulation and design, such as the effectiveness of monetary policy. In an attempt to evaluate one aspect of the efficacy of monetary policy in Jordan, the neutrality of money hypothesis is the focus of this paper.

The paper is organized as follows. The next section reviews the most recent studies conducted about the neutrality of money in different countries. The third section presents the econometric framework used to examine the neutrality issue. The empirical results are discussed in the fourth section. Section five concludes.

2. Literature Review

Much of the published literature in the last thirty years about the effects of short-term variations in money growth on output has emphasized a distinction between anticipated and unanticipated components of money growth (Frydman and Rappaport 1987). The empirical evidence based on this issue has largely been limited to an evaluation of two competing models. First, the New Classical model in which only unanticipated money growth over a short time horizon has real effects, and the so called Keynesians' model which presumes a real effect of actual money growth, whether unanticipated, or not (Barro 1977 and 1978). The New Classical model predicts that temporary and sustained changes will have the same effect. The sustained changes lose their effectiveness as soon as they are announced and expected. In contrast, the Keynesian model predicts that sustained changes will have different effects than temporary changes, and that sustained changes in the growth rate of money will not be neutral in the long run. Although the econometric literature on the relevance of this distinction is extensive, it remains inconclusive.

A recent study by Alkhatib (1995b) on the Of the most related study is the only empirical work done on the link between money and output. only empirical study of related tudy about the link between money and output has been published in Alkhatib (1995) examined the link between money supply and output in Jordan for the sample period from 1969 to 1991 on the basis of implementation of a final prediction approach. The empirical results revealed strong evidence supporting the hypothesis that money Granger-causes output, while the hypothesis that money reacts passively to output is rejected.

Hine and Bischoff (1998) suggested an alternative model, which is superior to the New Classical and Keynesian models. It is based on the ideas of Fischer (1977), who argues that errors in expectations of monetary growth extending over several periods are relevant to output determination due to the existence of long-term nominal wage contracts. This model treats fluctuations in output as a function of the difference between the growth rate of money supply and expectations of this rate formed as much as two years earlier. On the basis of this framework, expectational errors made six and seven quarters before the end of the period, over which money growth is measured are found to be important for output. This gives the monetary authority, enough time to react to events which occur after these expectations were formulated. The results thus provide strong empirical evidence for the role of long-term contracts in the determination of real activity, a role which has been recognized as being of great significance in the theoretical literature (Barro and Hercowitz 1980). These findings also provide a possible explanation for the failure of much of the related empirical research to reach an unambiguous conclusion (Azariadis, Bullard, and Ohanian 1999)

The Hine-Bischoff model predicts that a sustained increase in money, either its level or its growth rate, will, in the presence of rational expectations, have a substantially greater impact on output than a temporary increase of equal magnitude. For an unanticipated shock in money growth at time zero there exists a unique post shock equilibrium price, assuming some restrictions on shock size and parameter values. For a positive shock, the growth rate of prices initially rises above its long-run steady state rate. It then fluctuates about the steady state rate in damping cycles over time.

An empirical study conducted by Shelley and Wallace (1998) investigated the neutrality of anticipated and unanticipated money using U.S. disaggregated data for the period 1954:1 to 1994:12. The sample consists of the output of twenty manufacturing industries at the two digit Standard Industrial Classification (SIC) level. Seasonally adjusted industrial production indexes are used as the output measures and seasonally adjusted M1 as the money measure. The null hypothesis of neutrality is rejected for twelve of the considered fourteen industrial output series. In eight of these cases, neutrality is rejected at the 1% significance level. Neutrality of anticipated money fails to be rejected only for two industries. A similar approach is followed to test for the neutrality of unanticipated money. The hypothesis of unanticipated money neutrality is rejected at 5% significance level or better in seven of fourteen series.

Leong and Maki (2000) considered the long-run neutrality hypothesis using Australian data. A reduced form autoregressive integrated moving average (ARIMA) model is used with both quarterly seasonally unadjusted and adjusted Australian real GDP and nominal money supply to test the neutrality hypothesis. Using two measures of money stock; namely M1 and M3, it is shown that the hypothesis is supported using M1 as a measure of money supply, that is, changes in M1 have no effects on changes in real output. However, the long-run neutrality hypothesis is rejected using M3, in that changes in M3 significantly affect changes in real output. These results for Australia indicate the sensitivity of the outcome to the type of money supply used.

The long-run neutrality of money on real output was examined by Wallace (1999) for Mexico using a model developed by Fisher and Seater (1993) for the 1932-92 period. The results of the estimation indicate that changes in the quantity of money have no effect on the level of real output in the long run. The conclusion is robust whether M1 or M2 is used as a money measure. It is also robust for an alternative specification with a time period dummy for the period during which domestic banks in Mexico were nationalized. This supports the long-run neutrality proposition, which is noteworthy in the context of Mexico's stormy banking and monetary history during the study period.

Fisher and Seater (1993) found that long-run neutrality was rejected by the US annual data from 1869 to 1975. The seminal research by Fisher and Seater, which uses a non-structural approach to test the neutrality hypothesis, has led to numerous related publications. For example, Bullard (1994) and Serletis and Krause (1996) adopted Fisher and Seater's approach using the US data and found that the long-run neutrality hypotheses were generally supported. Boschen and Otrok (1994) and Olekalns (1996) adopted Fisher and Seater's approach to analyse the money supply-output relationship in the USA and Australia, respectively, accommodating structural breaks with split samples and dummy variables. It was found that the outcomes of the tests were not robust to structural breaks. Recently, Haug and Lucas (1997) successfully used Canadian data to verify the Boschen and Otrok finding that structural breaks can adversely affect the outcomes of neutrality tests.

A recent paper by Coe and Nason (1999) also contributes to this literature. They used the Fisher and Seater (1993) test for long-run neutrality, and they used the same U.S. data as Fisher and Seater, except that they update the data through 1997. When Coe and

Nason use a broad measure of the money stock (as Fisher and Seater did), they supported the Fisher and Seater rejection of long-run

monetary neutrality with respect to real output. But when they replaced the broad money measure by the monetary base, they could no

longer reject long-run neutrality. They also considered data from the United Kingdom for about a century and fail to reject long-run neutrality using either broad or narrow measures of money. Coe and Nason have concluded that the Fisher and Seater rejection of long-run neutrality is not robust to different measures of money or different countries.

Serletis and Krause (1996) have used a data set that includes more than 100 years of annual observations on real output, prices, and money for Australia, Canada, Denmark, Germany, Italy, Japan, Norway, Sweden, the United Kingdom, and the United States. They utilized the unit roots test using the procedures of Zivot and Andrews (1992), and they conclude that money is reasonably described as $I(1)$ except in Germany and Japan where it was $I(0)$; these latter two countries are therefore uninformative on neutrality questions in this data set. Serletis and Krause (1996) have found that output is $I(0)$ for Australia, Canada, Denmark, Italy, the United Kingdom, and the United States. These countries, therefore, provide direct evidence in favor of the longrun neutrality with respect to output. Serletis and Krause use the Fisher and Seater (1993) regression test to produce estimates for the remaining money-price or money-output combinations. Their results generally support the hypothesis of long-run monetary neutrality.

3. The Econometric Framework

The dynamic relationships between two bivariate time series, (m_t and y_t) can be expressed as

$$\begin{pmatrix} m_t \\ y_t \end{pmatrix} = \begin{pmatrix} F_{11}(B) & F_{12}(B) \\ F_{21}(B) & F_{22}(B) \end{pmatrix} \begin{pmatrix} e_{1t} \\ e_{2t} \end{pmatrix} \quad (1)$$

where e_{1t} and e_{2t} are independent white noise innovations and the $F_{ij}(B)$ are convergent rational functions in B , the backshift operator ($B^k m_t = m_{t-k}$), the focalized variables m_t and y_t are money supply and real output, respectively.

Following an approach developed by Leong and McAleer (2000), it is assumed that the relationship between money supply and real output can be represented by a stationary and invertible bivariate log-linear ARIMA model. It is important to use natural logarithmic transformations of variables because first differences approximate percentage changes, which are valuable for an analysis of neutrality. The model is given as follows:

$$\begin{pmatrix} m_t \\ y_t \end{pmatrix} = \begin{pmatrix} \Psi_{11}(B) & 0 \\ 0 & \Psi_{22}(B) \end{pmatrix} \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix} \quad (2)$$

where $\Psi_{ii}(B)$ are also convergent rational function in B . Now if m_t and y_t are related through $F_{12}(B)$ or $F_{21}(B)$, then ε_{1t} and ε_{2t} are also related. An appropriate way to examine the dynamic relationship between m_t and y_t is to estimate the cross-correlation function between ε_{1t} and ε_{2t} . The cross-correlation function between ε_{1t} and ε_{2t} is defined as

$$\rho_{\varepsilon_1 \varepsilon_2}(k) = \frac{\gamma_{\varepsilon_1 \varepsilon_2}(k)}{\sigma_{\varepsilon_1} \sigma_{\varepsilon_2}} \quad k = 0, \pm 1, \pm 2, \dots \quad (3)$$

where $\gamma_{\varepsilon_1 \varepsilon_2}(k) = E(\varepsilon_{1t-k} \varepsilon_{2t})$, $\sigma_{\varepsilon_1}^2 = E(\varepsilon_{1t}^2)$, $\sigma_{\varepsilon_2}^2 = E(\varepsilon_{2t}^2)$. Given n observations on ε_{1t} and ε_{2t} , $\rho(k)$ can be estimated by

$$r_{\varepsilon_1 \varepsilon_2}(k) = \left[\sum_{t=1}^n \varepsilon_{1t-k}^2 \sum_{t=1}^n \varepsilon_{2t}^2 \right]^{0.5} \sum_{t=k}^n \varepsilon_{1t-k} \varepsilon_{2t} \dots \dots \dots (4)$$

$$k = 0, \pm 1, \pm 2, \dots, \pm M$$

On the basis of this correlation function, Haugh (1976) has developed a popular framework for this analysis using the univariate residual cross-correlation function. This function reveals the nature of any empirical relationship between the two series in question and provides a method for testing the hypothesis that they are independent. The statistic is

$$S = n \sum_{k=-M}^M r_{\varepsilon_1 \varepsilon_2}(k)^2 \quad \text{for large samples} \quad (5)$$

or

$$S^* = n \sum_{k=-M}^M (n-|k|)^{-1} r_{\varepsilon_1 \varepsilon_2}(k)^2 \quad \text{for small samples} \quad (6)$$

Under the null hypothesis, S is asymptotically distributed chi square with $(2M + 1)$ degrees of freedom. The hypothesis of independence will be rejected in the presence of a number of relatively large cross-correlation coefficients. The main deficiency of this test is that it ignores any potential pattern in successive coefficients in the cross-correlation function. In more specific words, the Haugh statistic is composed of the sum of squared coefficients, so the successive positive or negative coefficients that are arranged in a distinct pattern are given the same weights.

Koch and Yang (1986) suggested a formal statistic r^* , in which information is incorporated about a possible pattern among successive coefficients. The strength of this test is that it distinguishes one cross-correlation function with estimated coefficients that are small in magnitude and randomly distributed about zero from another cross-correlation function coefficients that are small in magnitude but arranged in a distinct pattern. This statistic is defined as

$$r_i^* = n \sum_{k=-M}^{M-1} \left\{ \sum_{l=0}^i r_{k+l}^2 \right\} \quad (7)$$

$i = 0, 1, 2, \dots, M-1$

where r_i^* is approximately distributed as $\beta_i \chi_{\eta_i}^2$. The parameter, β_i and η_i , depend on the moments of r_i^* as follows:

$$\beta_i = \frac{\text{tr}(A_i A_i)}{\text{tr}(A_i)}; \eta_i = \frac{\text{tr}(A_i)^2}{\text{tr}(A_i A_i)} \quad (8)$$

Thus $r_i^* \approx \beta_i \chi_{\eta_i}^2$ or $(1/\beta_i) r_i^* \approx \chi_{\eta_i}^2$, where β_i and η_i are given in (8). As shown by Koch and Yang, this approximate distribution is simple to apply. The individual eigenvalues of A_i need not to be calculated. Only the sum of eigenvalues and the sum of squared eigenvalues are necessary. Furthermore, the nature of matrix A_i allows straightforward calculation of these sums, as follows:

$$\text{where } \text{tr}(A_i) = 2(M+1)(i+1) - (i+1)^2 \quad (9)$$

$$\text{And } \text{tr}(A_i A_i) = 2 \sum [2M+i-3(j-1)]^2 + (2M+1-2i)(i+1)^2 \quad (10)$$

$$i = 0, 1, 2, \dots, M-1$$

This study will use two measures for money, (M1) and (M2), to investigate the robustness of results across different money measures. The industrial production index (1997 = 100) is used for the economic activity. These aggregates are not seasonally adjusted, and measured at the end of the month. The proposed r_i statistics from (7) will be calculated for three lag lengths (24, 30, and 36 months).

4. The Empirical Results

This section proceeds by firstly employing the autoregressive integrated moving average (ARIMA) to filter the original series to white noise. Since the theory behind ARIMA estimation applies only to cointegrated time series, the Augmented Dickey-Fuller test will be employed to test the stationarity of

the time series used in this context. The failure to properly transform nonstationary data into stationary data can result in model misspecification thus leading to incorrect inferences (Enders 1995). This test consists of running a regression of the first difference of the series (X_t) against the series lagged once, and k lagged difference terms

$$\Delta X_t = \beta_0 X_{t-1} + \beta_1 \Delta X_{t-1} + \beta_2 \Delta X_{t-2} + \dots + \beta_k \Delta X_{t-k} + e_t$$

In each case the results of the ADF test consists of the t-statistic on the coefficient of the lagged variables β_0 , and critical values of the test of a zero coefficient. If the coefficient is significantly different from zero then the hypothesis that X_t contains a unit root is rejected and the hypothesis that X_t is stationary is not rejected (Stock 1990, and Rudebusch 1993).

With six lagged difference terms, the ADF test was carried out for each series after taking the first, and twelfth differences of the natural log of these series. Results of ADF test are reported in Table 1. The Dickey-Fuller t-statistic is greater (in absolute value) than the reported critical value (2.58) at 1% level of significance, rejecting the existence of a unit root hypothesis of stationarity (Christiano and Eichenbaum 1990). These results provide empirical evidence showing that each series is integrated of order one $I(1)$. A series is $I(1)$ if its first difference does not contain a unit root.

Table (1)

Augmented Dickey-Fuller Unit Root Test Based on six lag lengths.

Explanatory Variables	$(1 - B)^2(1 - B)\log(y)$		$(1 - B)^2(1 - B)\log(M_0)$		$(1 - B)^2(1 - B)\log(M_1)$	
	Coefficient	t-Statistic	Coefficient	t-Statistic	Coefficient	t-Statistic
L(-1)						
$\Delta L(-1)$	-2.665	-5.964	-1.837	-4.580	-0.728	-4.391
$\Delta L(-2)$	1.130	2.769	0.298	0.794	-0.223	-1.409
$\Delta L(-3)$	0.772	2.195	0.002	0.005	-0.005	-0.030
$\Delta L(-4)$	0.527	1.819	-0.096	-0.331	0.064	0.449
$\Delta L(-5)$	0.261	1.1452	-0.112	-0.485	0.097	0.738
$\Delta L(-6)$	0.035	0.218	-0.037	-0.224	0.200	1.728
$\Delta L(-6)$	0.012	0.148	0.027	0.303	0.087	1.03
ADF Test Statistic		-5.963**		-4.580**		-4.391**
SE of Regression		0.092		0.060		0.501
Number of Observations		132		132		132
F-Statistic		55.169		56.200		22.894
R ²		0.726		0.730		0.524
Akaike Information Criterion		-4.718		-5.574		-8.407

** The critical value at 1% significance level is 2.58

L(-1) refers to the lag 1 of the variable, and Δ refers to the change

The second step in the analysis is to employ the Box-Jenkins cycle of preliminary identification, estimation, and diagnostic checking to filter the stationary series to white noise. The chosen univariate ARIMA models are given in table 2. These univariate ARIMA are used to filter the original time series, Y, M1 and M2 to ε_1 , ε_2 , and ε_3 , respectively. The cross-correlation functions are calculated relating the empirical distributed lags under scrutiny for 36 months in each direction (see Appendix A).

Asymptotic standard errors (Se) are given in parentheses in column four. The residual standard error for each model is calculated as the square root of the residual sum of squares divided by the number of residuals minus the number of free parameters in the model. Inspection of the Ljung-Box (1978) statistic, Q, indicates that these filters reduce their respective series to white noise. The marginal significance level of the Ljung-Box statistic appears in parentheses in the fifth column of table (2).

Table (2)

Estimated Univariate ARIMA models

Variable	Coefficient Estimates	Standard error	Q -statistic ^a	DF ^b	SE of Regression	
(1-B)(1-B ¹²)logM _{1t}	AR(1)	-0.5530	(0.0659)	24.712 (0.479)	25	0.0507
	AR(2)	-0.4048	(0.0741)			
	MA(3)	-0.1991	(0.0450)			
	MA(10)	-0.3603	(0.0463)			
	MA(12)	-0.3506	(0.0469)			
(1-B)(1-B ¹²)logM _{2t}	AR(9)	-0.1434	(0.0682)	28.395 (0.444)	28	0.0095
	MA(12)	-0.8857	(0.0001)			
(1-B)(1-B ¹²)logy _{1t}	AR(1)	-0.6213	(0.0714)	27.910 (0.264)	24	0.0694
	AR(2)	-0.4552	(0.0816)			
	AR(3)	-0.2859	(0.0847)			
	AR(4)	-0.2550	(0.0811)			
	AR(5)	-0.1894	(0.0741)			
	MA(12)	-0.8857	(0.0001)			

$$^a Q = n(n+2) \sum_{k=1}^M (r_{\alpha k}^2 / n-k)$$

Where $r_{\alpha k}$ is the k-th residual autocorrelation, n is the number of observations, M is the number of the residual autocorrelations if the series has not previously been subject to ARIMA analysis. Q is approximately distributed as $\chi^2_{(M-p-q)}$, p and q are the orders of the autoregressive and the moving average process, respectively.

^b DF the degree of freedom is the number of residual autocorrelations less the number of autoregressive and moving average terms previously estimated.

The third step in the analysis is to conduct the neutrality test using Koch-Yang statistic, r_i^* , which has been computed for all alternative cross-correlation functions for three lag lengths: $M = 36, 30,$ and 24 months. The neutrality hypothesis between any two pairs of cross-correlation functions will be rejected if the corresponding statistic $(1/\beta_i) r_i^*$, exceeds the critical $\chi^2_{\eta_i}$ at the 5% significance level.

First, we consider the independence hypothesis between money supply M1 and industrial production index. The results appear in table 3. It is very clear that the neutrality hypothesis is not rejected at the 5% significance level for all three lag lengths. These results are consistent with the cross-correlation function coefficients reported in appendix A. The cross-correlation function of series M1 and industrial production displays no significant spikes at any usual significance level, revealing empirical evidence supporting the neutrality hypothesis of M1 in both short-run and long-run as well. However, the function displays two weak distinct patterns of cross-correlation at lags $-24, \dots, -29$ (of positive values), and at $-30, \dots, -36$ (of negative values). These two patterns may not necessarily imply a substantive relationship between growth rate of M1 and the growth rate of industrial production index.

Table (3)
Koch-Yang Test results for Testing the Independence of Money Supply (M1) and Index of Industrial Production (IIP).

i	M=24			M=30			M=36		
	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α
0	49.0	25.90	1.00	61.0	36.31	1.00	73.0	41.82	1.00
1	32.2	16.95	0.99	40.2	26.95	0.94	48.2	31.88	0.97
2	22.6	13.28	0.93	28.2	22.60	0.75	33.9	26.07	0.80
3	17.1	10.17	0.90	21.4	18.99	0.59	25.8	22.02	0.64
4	13.6	8.42	0.82	17.1	16.37	0.50	20.6	18.97	0.52
5	11.2	6.84	0.81	14.2	13.74	0.47	17.1	16.22	0.51
6	9.5	5.82	0.76	12.0	11.76	0.47	14.6	14.23	0.43
7	8.2	4.85	0.77	10.4	9.98	0.44	12.7	12.13	0.44
8	7.2	4.23	0.75	9.2	8.67	0.47	11.1	10.53	0.49
9	6.4	3.60	0.73	8.1	7.40	0.49	9.9	9.25	0.42
10	5.7	3.25	0.66	7.3	6.62	0.47	8.9	8.24	0.41
11	5.1	2.78	0.73	6.6	5.75	0.45	8.1	7.35	0.50
12	4.7	2.41	0.66	6.0	5.18	0.52	7.4	6.64	0.47
13	4.3	2.09	0.72	5.5	4.60	0.47	6.8	5.92	0.43
14	3.9	1.87	0.60	5.1	4.11	0.53	6.3	5.29	0.51
15	3.6	1.71	0.64	4.7	3.64	0.46	5.8	4.77	0.44
16	3.3	1.57	0.67	4.4	3.23	0.52	5.4	4.29	0.51
17	3.1	1.47	0.69	4.1	2.78	0.60	5.1	3.85	0.57
18	2.9	1.35	0.51	3.8	2.36	0.50	4.7	3.44	0.49
19	2.5	1.22	0.54	3.6	2.01	0.57	4.5	3.01	0.56
20	2.4	1.10	0.58	3.4	1.75	0.63	4.2	2.64	0.62
21	2.2	0.99	0.61	3.2	1.48	0.69	4.0	2.30	0.68

i	M = 24			M = 30			M = 36		
	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α
22	2.1	0.88	0.65	3.0	1.29	0.73	3.7	2.03	0.57
23	2.0	0.73	0.69	2.8	1.11	0.58	3.5	1.74	0.63
24				2.7	0.94	0.63	3.4	1.48	0.69
25				2.5	0.86	0.65	3.2	1.30	0.73
26				2.4	0.79	0.68	3.0	1.16	0.76
27				2.2	0.71	0.70	2.8	0.98	0.61
28				2.1	5.69	0.71	2.6	0.91	0.64
29				2.0	0.72	0.70	2.5	0.89	0.64
30							2.4	0.89	0.64
31							2.3	0.91	0.64
32							2.2	0.94	0.63
33							2.2	0.97	0.62
34							2.1	0.99	0.61
35							2.0	1.02	0.60

According to the results of Koch-Yang test based on the cross-correlation function between money supply, M2, and industrial production, we cannot reject the neutrality hypothesis at any usual significance level. This independence is basically due to the fact that the cross correlation function between these two series displays only three significant spikes at lags -11, 10, and 11. These spikes were not enough to reject the null hypothesis of independence, especially no distinct patterns of coefficients are observed.

Table (4)

Koch-Yang Test results for Testing the Independence of Money Supply (M2) and Index of Industrial Production (IIP).

i	M = 24			M = 30			M = 36		
	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α
0	49.0	53.72	0.30	61	64.93	0.34	73.0	67.44	0.66
1	32.2	23.07	0.88	40.2	33.04	0.77	48.2	34.73	0.92
2	22.6	15.88	0.82	28.2	23.59	0.70	33.9	24.79	0.85
3	17.1	12.06	0.80	21.4	17.80	0.66	25.8	18.86	0.80
4	13.6	8.61	0.80	17.1	12.13	0.79	20.6	13.64	0.85
5	11.2	6.34	0.30	14.2	8.70	0.85	17.1	10.45	0.88
6	9.5	5.24	0.81	12.0	7.01	0.86	14.6	8.82	0.84
7	8.2	4.33	0.83	10.4	5.80	0.83	12.7	7.55	0.82
8	7.2	3.68	0.82	9.2	5.11	0.82	11.1	6.66	0.83
9	6.4	3.46	0.75	8.1	4.79	0.78	9.9	6.27	0.71
10	5.7	2.86	0.72	7.3	4.24	0.75	8.9	5.41	0.71
11	5.1	2.25	0.81	6.6	3.50	0.74	8.1	4.48	0.81
12	4.7	1.89	0.76	6.0	3.13	0.79	7.4	3.97	0.78
13	4.3	1.45	0.84	5.5	2.54	0.77	6.8	3.42	0.76
14	3.9	1.12	0.77	5.1	2.35	0.80	6.3	3.23	0.78
15	3.6	0.83	0.84	4.7	2.24	0.69	5.8	3.13	0.68
16	3.3	0.63	0.89	4.4	2.10	0.72	5.4	2.97	0.71
17	3.1	0.53	0.91	4.1	2.02	0.73	5.1	2.86	0.72
18	2.9	0.41	0.82	3.8	1.83	0.61	4.7	2.72	0.61
19	2.5	0.29	0.87	3.6	1.61	0.66	4.5	2.46	0.65
20	2.4	0.27	0.88	3.4	1.33	0.72	4.2	2.32	0.68
21	2.2	0.28	0.87	3.2	1.10	0.78	4.0	2.21	0.70

i	M = 24			M = 30			M = 36		
	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α	η_i	$(1/\beta_i)r_i^*$	α
22	2.1	0.27	0.88	3.0	0.92	0.83	3.7	2.09	0.56
23	2.0	0.29	0.87	2.8	0.83	0.66	3.5	1.97	0.58
24				2.7	0.69	0.71	3.4	1.77	0.62
25				2.5	0.57	0.75	3.2	1.55	0.67
26				2.4	0.54	0.77	3.0	1.32	0.72
27				2.2	0.47	0.18	2.8	1.04	0.60
28				2.1	0.49	0.78	2.6	0.93	0.63
29				2.0	0.50	0.78	2.5	0.86	0.65
30							2.4	0.81	0.67
31							2.3	0.76	0.69
32							2.2	0.73	0.70
33							2.2	0.69	0.71
34							2.1	0.67	0.72
35							2.0	0.68	0.71

Using two measures of money stock; namely M1 and M2, it is shown that the hypothesis of neutrality is not rejected for all lag lengths used. Furthermore, the response patterns of industrial production to money seem to be similar when different money measures are used. The neutrality of money during this period may be attributed to the developments in the money and credit markets in Jordan. During this period, the Central Bank of Jordan intensified its indirect intervention in the money market to accommodate any potential possible increase in the demand for foreign currencies. The Monetary policy was conducted in a way to maintain a suitable level of foreign reserves to ensure stability for the exchange rate of the Jordanian dinar without jeopardizing its convertibility. On the other hand, the neutrality could be attributed to the methodology employed or the data set used. The use of industrial production index as a measure of economic activity is of limited usefulness for drawing substantive economic relations.

5. Concluding Remarks

We have investigated the dynamic relationships between money and industrial production index by using Jordan monthly data over a period starting from January 1990 through August 2002. The estimated cross-correlation functions and the formal statistic suggested by Koch and Yang provide no evidence against the neutrality of money in both short-run and long-run. These conclusions are robust to three-lag lengths (24, 30, and 36 months), and two different money measures (M1 and M2).

The sign and magnitude of the estimated effect of money growth on the level of output depend critically on the specific identifying restriction employed. The results presented above should be interpreted after taking account of the following caveats. First, the results are predicated on specific assumptions concerning the degree of integration of the data, and with 140 observations for monthly data the degree of integration probably uncertain.

Second, the analysis has been carried out using bivariate models. If there are more than two important sources of macroeconomic shocks, then bivariate models may be subject to significant omitted variable bias. Thus another extension of this work is to expand the set of variables under study to allow a richer set of structural macroeconomic shocks. Third, real output, in this context, is measured by the index of industrial production index. It may be worth noting that this measure is not necessarily the best measure for economic activity. However, it may be the best among all the variables that are published on monthly basis.

الحيادية النقدية في الأجل الطويل في الأردن

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ملخص

قامت الدراسة باختبار فرضية حيادية النقود في الأجل الطويل في الأردن باستخدام بيانات شهرية خلال الفترة الزمنية من كانون الثاني 1990 وحتى آب 2003. استخدمت الدراسة مقياسين لعرض النقد (ع1) و (ع2) لقياس المؤشرات النقدية، و الرقم القياسي للإنتاج الصناعي (1997 = 100) كمقياس للنشاط الاقتصادي. وقد تم استخدام اختبار كوك - بانغ لتقصي العلاقة المحتملة بين هذه المتغيرات. بينت النتائج أنه لا يوجد أي علاقة ذات دلالة إحصائية بين مقياسي عرض النقد (ع1) و (ع2) من ناحية وبين الرقم القياسي للإنتاج الصناعي خلال هذه الفترة من ناحية أخرى. وكانت النتائج متشابهة إلى حد بعيد وذلك على الرغم من اختلاف أطوال الفترات التي تم على أساسها اختبار العلاقة.

Appendix A

Univariate Cross-Correlation Functions for Money Series, M1, M2, and Industrial Production Index (IIP).

K	S(r)	r_{e1i}	r_{e2i}	K	S(r)	r_{e1i}	r_{e2i}
-36	0.102	-0.105	-0.062	0	0.087	0.065	0.012
-35	0.102	-0.068	-0.039	1	0.087	0.105	0.107
-34	0.101	-0.009	0.009	2	0.088	0.046	0.045
-33	0.101	-0.011	0.084	3	0.088	0.073	0.020
-32	0.100	-0.005	-0.042	4	0.088	-0.015	0.003
-31	0.100	-0.035	0.039	5	0.089	-0.020	-0.070
-30	0.099	-0.058	-0.047	6	0.089	-0.096	-0.005
-29	0.099	0.093	-0.001	7	0.089	0.087	0.019
-28	0.098	0.097	0.173	8	0.090	-0.009	0.021
-27	0.098	0.071	0.142	9	0.090	0.006	-0.063
-26	0.097	0.160	0.017	10	0.091	0.107	0.278**
-25	0.097	0.087	0.029	11	0.091	0.054	-0.276*
-24	0.096	0.029	0.002	12	0.091	0.168	0.082
-23	0.096	-0.024	-0.039	13	0.092	0.036	-0.101
-22	0.095	0.094	0.066	14	0.092	-0.070	-0.063
-21	0.095	-0.071	-0.109	15	0.092	0.025	0.122
-20	0.094	0.059	0.096	16	0.093	-0.056	0.042
-19	0.094	-0.056	0.013	17	0.093	0.078	-0.013
-18	0.094	0.033	0.118	18	0.094	-0.009	-0.056
-17	0.093	0.027	-0.056	19	0.094	-0.022	-0.011
-16	0.093	0.061	0.059	20	0.094	-0.106	-0.110
-15	0.092	-0.041	-0.067	21	0.095	-0.134	-0.058
-14	0.092	0.010	0.072	22	0.095	-0.044	0.056
-13	0.092	0.023	0.121	23	0.096	-0.053	0.055
-12	0.091	-0.007	-0.060	24	0.096	-0.071	0.068
-11	0.091	-0.067	0.181*	25	0.097	0.010	0.008
-10	0.091	0.063	0.090	26	0.097	-0.103	-0.179
-9	0.090	-0.093	-0.139	27	0.098	-0.012	-0.011
-8	0.090	0.007	-0.046	28	0.098	-0.076	0.039
-7	0.089	-0.011	0.025	29	0.099	0.037	0.019
-6	0.089	-0.142	-0.123	30	0.099	0.070	0.056
-5	0.089	0.003	-0.054	31	0.100	-0.029	0.013
-4	0.088	0.065	-0.016	32	0.100	0.096	0.009
-3	0.088	-0.013	-0.105	33	0.101	0.131	-0.041
-2	0.088	-0.053	0.084	34	0.101	-0.016	-0.010
-1	0.087	-0.045	-0.100	35	0.102	0.000	-0.048
0	0.087	0.065	0.012	36	0.102	0.009	0.055

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