The Role of Capital Formation and Saving in Promoting Economic Growth in Iran

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ABSTRACT

This paper estimates the interdependencies between real capital formation, saving and output for Iran in the turbulent years 1960 to 2003. The analysis uses Zivot and Andrews (1992) procedure to endogenously determine that structural breaks occurred in 1984 for real output, 1980 for saving and 1979 for investment. These dates coincide with the effect of the Islamic revolution in 1979 and Iran-Iraq war from 1980 to 1988.

The Johansen FIML estimates for the non-stationary variables indicate significant Solow style relationships between output and saving in the long run, whilst investment is found to be imprecise. The long run share of income that is paid to capital (in the form of saving) is estimated at 0.55, which is higher than the 0.33 average for developed countries.

The short run estimates further support the Solow model whereby changes to saving have transitory equilibrating effects on the growth in output. However, investment dynamically Granger causes the short run growth in output, consistent with endogenous growth explanations. The structural change parameter estimates that the growth in output fell by around 10 percent after 1979.

These findings have two important policy implications for Iran. First, there is scope to reduce the reliance of saving as the domestic source of economic growth. Second, saving needs to be better targeted to the long run strategic provision of capital (including infrastructure) in the structurally transforming economy of Iran.

Keywords: Economic growth, saving, investment, cointegration.
JEL Classifications: C13, E21, O16

I Introduction

This paper investigates how capital formation and saving promote economic growth in Iran. This is a challenging task given the unresolved debate about the roles of investment and saving (both empirically and theoretically) in models of growth and the difficulty of specifying and estimating the relationships for an economy which has experienced profound changes over the past four decades. We believe it is necessary to briefly consider each of these important factors in turn.

Houtakker (1961, 1965), Modigliani (1970) and many others provide empirical evidence of the positive correlation between saving and output for a large number of countries. This direct relationship is often argued as supporting the Solow style model of growth in which a higher saving rate causes transitory growth to a higher steady state level of output. However there is growing evidence that causation may run in the other direction, from growth to saving, called the Carroll-Weil hypothesis.[1] There is further disagreement about the subsequent effect of saving on investment. Whilst Feldstein and Horioka (1980) emphasized the powerful empirical association between saving and investment, no consensus explanation has emerged about this link or its direction.

Levine and Renelt (1992) use cross–country data to show that investment is the only variable that is robustly correlated with the growth in output. Whilst most argue the causal link is from investment to output, there is some evidence that output influences investment through an accelerator effect. The possible complex feedback effects and observed variations in productivity are consistent with the endogenous growth view. Hall and Jones (1999) argue that most cross-sectional variation in per capita output is due to variation in the productivity with which factors are combined, rather than differences in factor accumulation. King and Levine (1994) provide evidence that capital accumulation alone is neither a necessary nor sufficient condition for the "take-off" to rapid growth.

These unresolved issues provide only broad guidance for researchers and policy makers, whose task is made even more difficult when studying developing countries with individual and specific characteristics like that of Iran. To the best of our knowledge, there

[1] This is most evident in the East Asian economies which had high growth rates long before they had high saving rates. Similarly, Japan had a high income growth in the late 1940s and early 1950s, yet Japan did not exhibit high saving rates until 1960s and 1970s.

are few studies which consider the effects of saving and investment on economic growth in the Middle East and even fewer for Iran. Eken, Helbling and Mazarei (1997) show that, for non-oil exporting countries, the share of private investment is positively correlated with economic growth in countries in the MENA region.

However Iran is a major oil exporter and Jalali-Naini (2003) claims the “basic development thinking in Iran since the mid 1950s has been a planning framework in which the oil industry, as the ‘leading sector’ and the engine of growth supplies surpluses (saving) for investment in other sectors”, (p. 18).[2] Indeed, government policies have been very important in Iran’s economic performance over the last four decades.[3] Table 1 shows that real gross domestic product (GDP), gross national saving (GNS) and gross fixed capital formation (GFCF) grew strongly and consistently from 1960 to 1978 in line with the growth in the private sector.[4] However the high co-movements in these variables ended when the sharp increase in crude oil prices in 1974 fuelled an economic boom, causing higher inflation which adversely affected economic growth in the late 1970s.

Table 1

Real GDP, Saving and Investment Growth Rates (percent)

Era	Period	Real GDP	Real Saving (GNS)	Real Investment (GFCF)
Pre- Revolution	1960-78	9.0	16.2	11.4
Post- Revolution	1979-03	2.5	6.2	4.3
- War years	1980-88	-1.5	6.5	-1.5
- First plan	1989-94	7.5	7.7	4.6
- Second plan	1995-99	3.2	5.2	10.1

Sources: National Accounts, Central Bank of Iran (2001), Hakimian (1999).

The boom ended with the Islamic revolution in 1979, which introduced significant changes to economic policies. There was extensive nationalization and greater state

2 He also finds that total factor productivity has not contributed to economic growth in Iran for the period 1959 to 2000.

3 Bahmani-Oskooee (1993) analyses the effects of official exchange control via the black market exchange rate effects on purchasing power parity.

4 The strong growth (although there was a dip in GNS in 1974-75) was due to a combination of low inflation, an increase in the demand for domestic money and a stable exchange rate.

control of prices in regard to large-scale modern industries, the banking and insurance sectors as well as foreign trade. Jalali-Naini (2003) notes that these policies, together with economic mismanagement and institutional and public sector inefficiency, caused high levels of uncertainty and resource misallocation.

Even more devastating to the economy was the eight-year war with Iraq, which assured that inappropriate government interventionist policies would continue. During the war years (1980-88) Iran experienced low investment and productivity with negative growth in output.[5] The physical damage of the war has been estimated to be around 30,811 billion Rials (Mazarei, 1996). Another adverse effect in this period occurred with the oil crisis in 1986 and the sharp drop in foreign exchange receipts from oil revenue in which led to the 1986-88 recession. According to Mazarei (1996), the difficulty in importing intermediate and capital goods due to the lack of foreign exchange was one of the causes of serious problems on the supply side of the Iranian economy at this time.

A new period of reconstruction began with the end of the war in late 1988 and economic adjustment policies were implemented under the First Five-Year Development Plan (FYDP). During 1989-1994, real GDP increased by 7.5 percent, while saving and investment increased by 7.7 and 4.6 percent respectively. Pesaran (2000) attributes this growth to the liberalization of trade and foreign exchange markets together with the utilization of previously unused capacity in the economy. Jalali-Naini (2003) refers to other relevant factors like the loosening of some government controls, partial correction of the prices system and a move towards privatization which were part of the government's 'structural adjustment policies'. Investment responded by increasing at a rate of 10.1 percent during the Second Five-Year Development Plan, 1995-99, whilst the growth in saving was only half of this at 5.2 percent.

This brief review of Iran's economy shows the difficulty in disentangling the complex and changing interrelationships between output, saving and investment for the period of our study, 1960 to 2003. It is essential that structural change in a growth setting is explicitly incorporated into the simultaneous analysis of these interdependencies. The next section therefore tests for structural change and non-stationarity in the variables, which are then

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[5] Direct war expenditures comprised on average 16.9 percent of total Iranian government expenditures between 1981 and 1986.

incorporated into the simultaneous estimation of their dependencies in the Section III. The final section summarises the key findings and brings out some policy implications.

II Unit Root Tests with Structural Breaks

It is well known that if potential structural breaks are not allowed for in testing for unit roots in time series, the tests may be biased towards a mistaken non-rejection of the non-stationarity hypothesis (Perron 1989, 1997; Leybourne and Newbold 2003; Pahlavani *et al*. 2005). Given Iran's experience it is surprising that very few studies of the Iranian economy has considered the issue of structural breaks. An exception is Bahmani-Oskooee (1993) who assumed a structural break occurred in 1979 when examining the effects of the black market exchange rate on relative prices.

Christiano (1992) and others have criticized of using a known exogenous structural break, arguing that this invalidates the distribution theory underlying conventional testing (Vogelsang and Perron, 1998). In response, a number of studies have proposed different ways of estimating the time of the break endogenously which lessen the bias in the usual unit root tests.[6] The null hypothesis in the Zivot and Andrews (1992) method is that the variable under investigation contains a unit root with a drift that excludes any structural break, while the alternative hypothesis is that the series is a trend stationary process with a one-time break occurring at an unknown point in time. In this approach, the time of break (T_b) is chosen to minimize the one-sided t-statistic for $\alpha = 1$.[7] The Zivot and Andrews model endogenises one structural break in the intercept and trend of a time series, y_t, according to the hypotheses:[8]

$$\text{H}_0\text{: } y_t = \mu + y_{t-1} + e_t \tag{1}$$

$$\text{H}_\text{A}\text{: } y_t = \mu + \theta DU_t(T_b) + \beta t + \gamma DT_t(T_b) + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + e_t \tag{2}$$

The null hypothesis is rejected if α is statistically significant. The time of the break, T_b, is endogenously determined by running the model sequentially (allowing for T_b to be any year within a five percent trimming region) and selecting the most significant t-ratio for α.

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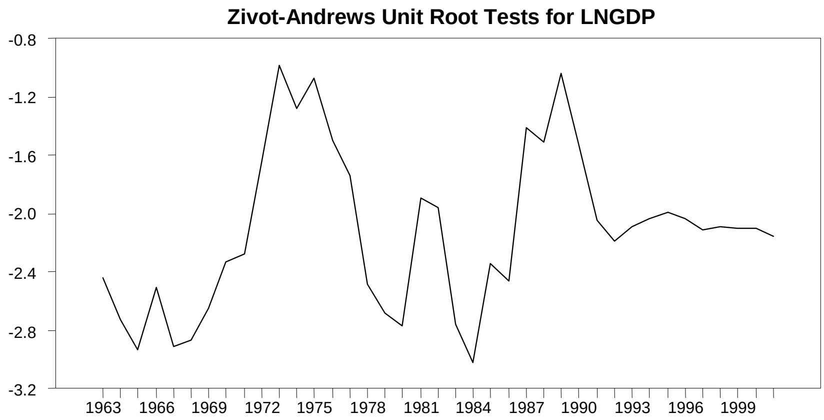
6 These studies include Perron (1994, 1997), Lumsdaine and Papell (1997) and Bai and Perron (2003).

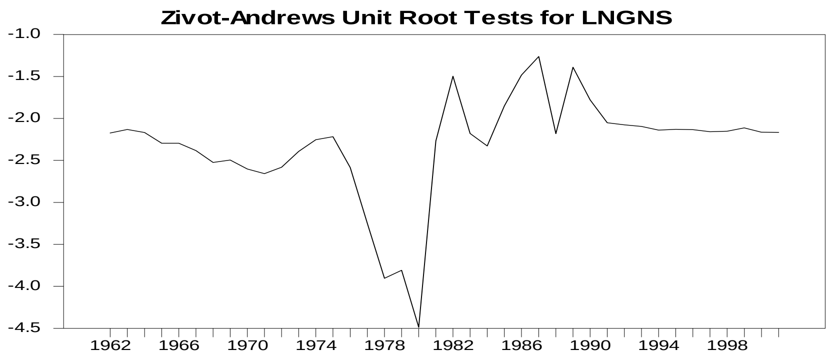
7 In other words, a break point is selected which is the least favourable to the null hypothesis.

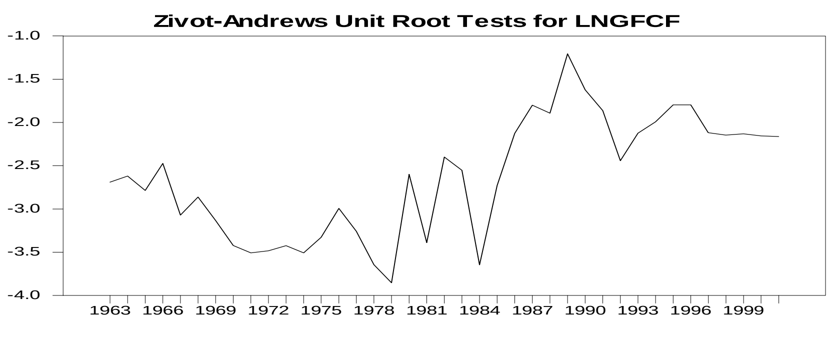
8 The dummy variable DU_t captures a shift in the intercept (where $DU_t = 1$ if $t > T_b$, and zero otherwise); DT_t captures a break in the trend occurring at time T_b (where DT_t is equal to (t - T_b) if ($t > T_b$) and zero otherwise). The computations were done using the RATS program.

Figure 1 plots the t_α values on the vertical axis and Table 2 summarises the test results for the sample period 1960 to 2003.

Figure 1







Note: The vertical axis shows the t_α values.
Source: National Accounts, Central Bank of Iran (2001; 2004).

Table 2

The Zivot-Andrews Test Results: Break in Both Intercept and Trend

	Symbol	k	T_b	θ	γ	α
Real GDP lnGDP)	Y	1	1984	–0.072 (–4.95)	–0.009 (–3.32)	–0.277 (–3.84)
Real Gross National Saving (lnGNS)	S	0	1980	–0.341 (–4.44)	–0.02 (–2.45)	–0.559 (–4.48)
Real Investment (lnGFCF)	I	1	1979	–0.152 (–3.09)	–0.024 (–3.02)	–0.471 (–3.85)

Notes: Critical values at 1, 5 and 10 percent levels are -5.57, -5.08 and -4.82, respectively (Zivot and Andrews, 1992).
The optimal lag length (k) is determined by the SBC.
Source: Central Bank of Iran (2001; 2004).

The Zivot and Andrews test results reveal that all of the variables under investigation are non-stationary, $I(1)$ for the estimated values of α. Table 2 also shows that the estimated coefficients for θ and γ are statistically significant, supporting the view that at least one structural break in the intercept and slope have occurred during the sample period for each variable. The time of the most significant structural break (T_b) is 1984 for real GDP, 1980 for GNS and 1979 for GFCF. It is interesting to note that the structural breaks in these variables coincide with major real events including the Islamic Revolution of 1979 and the beginning of the Iran-Iraq war in 1980. Because of the closeness of these years, we will select the start year of 1979 as the representative break date.

III Estimation of the Relationships

In order to test for the interdependent effects that the variables have on each other, it is necessary to use the Johansen's (1991, 1995) method.[9] The procedure is appropriate because it includes the specification and estimation of the simultaneous effects between the non-stationary variables. The VAR for the vector of variables, $\underline{X}_t{}' = \{Y_t, S_t, I_t\}$ is:

$$\underline{X}_t = \underline{\kappa} + \sum_{i=1}^{l} \Phi_i \underline{X}_{t-i} + \delta D_t + \upsilon_t \ , \qquad t = 1, 2, ..., n \tag{3}$$

with unrestricted intercepts $\underline{\kappa}$ and D_t the I(0) dummy variable taking value for 1979 to 2003 and zero otherwise.

9 See also Johansen and Julius (1992) and Pesaran and Pesaran (1997).

The model was estimated over the sample period, 1960 to 2003 for the optimum lag length, l, within the range of one to four lags. The model selection criteria and test statistics reported in Table 3 show possible optimum lags of 1, 2 and 3. Whilst there is supporting evidence of a lag of one according to the Schwarz Bayesian Criterion (SBC) and the Adjusted Likelihood Ratio (LR) test, it was decided to accept the lag of two since it is in the middle of the possible range, consistent with the Akaike Information Criterion (AIC) and allows for testing of Granger causality using the VECM (with reduced lag, $l-1=1$).

Table 3

Selection of the Optimum Lag Length (l)

Lag (l)	*AIC*	*SBC*	*LR Test*	*Adjusted LR*
4	87.71	52.25	–	–
3	87.08	59.21	19.28***	12.53
2	87.77	67.50	35.89	23.33
1	86.20	73.53	57.03	37.07**
0	3.39	–1.68	240.65	156.42

Notes: *AIC* represents the Akaike Information Criterion: *SBC* represents the Schwarz Bayesian Criterion: *LR* represents the Likelihood Ratio test:
*** Significant at the 1 percent level: ** 5 percent level: * 10 percent level.

The first order cointegrating VAR (with unrestricted intercept and no trend) gives the estimated eigenvalues: $\lambda_1 = 0.3552$, $\lambda_2 = 0.2418$, $\lambda_3 = 0.0906$ for $\Pi = \sum_{i=1}^{K} \Phi_i - \mathbf{I}$ having possible rank, $0 \le r \le 3$.[10] The smallest eigenvalue is close to zero and so the rank must be a maximum of two. However, the remaining values are also low, allowing the possibility of a rank of zero. The Likelihood Ratio tests and model selection criteria are shown in Table 4.

The maximal eigenvalue and trace tests accept the null hypothesis of $r=1$ and $r=2$ respectively, at the five percent levels of significance. All of the model selection criteria indicate a maximum rank of 3 which implies the system of three non-stationary variables is jointly stationary. It is likely that the lack of the degrees of freedom is affecting these criteria, which have relatively flat surfaces over the higher ranks. Since $r = 0$ implies no cointegration between the variables, it is sensible to not reject the null hypothesis, H_0: $r \le 1$ according to the *LR* test based on the trace of the stochastic matrix. This is consistent with selecting the largest of the (low) estimated eigenvalues, $\lambda_1 = 0.3552$.

Table 4

Selection of the Optimum Rank (*r*) of the Π Matrix
Likelihood Ratio (LR) Tests [1]

H_0	H_A	*Max λ*	H_A	*Trace*
$r = 0$	$r = 1$	18.43**	$r \ge 1$	34.05
$r \le 1$	$r = 2$	11.63	$r \ge 2$	15.62**
$r \le 2$	$r = 3$	3.99	$r = 3$	3.99

Model Selection Criteria [2]

Rank	*Max LL*	*AIC*	*SBC*	*HQC*
$r = 0$	100.71	85.70	72.67	80.93
$r = 1$	109.92	89.92	72.54	83.55
$r = 2$	115.73	92.73	72.75	85.41
$r = 3$	117.73	93.73	72.88	86.09

Notes: [1] *Max λ* represents the LR test based on the maximal eigenvalue of the stochastic matrix: *Trace* represents the LR test based on the trace of the stochastic matrix: *** Significant at the 1 percent level: ** 5 percent level: * 10 percent level.
[2] *Max LL* represents the maximum log of the likelihood function: *AIC* represents the Akaike Information Criterion: *SBC* represents the Schwarz Bayesian Criterion: *HQC* represents the Hann-Quinn Criterion.

The benefit of a rank of one is that we have only one cointegrating vector, $\underline{\beta}'\underline{X}_t$ from the decomposition, $\Pi = \underline{\alpha}\underline{\beta}'$. This reduces the required number of identifying restrictions on the cointegrating vector, $\{\beta_Y Y + \beta_S S + \beta_I I\} \sim \mathrm{I}(0)$ to a simple, single normalisation.[11] This is sufficient to identify the long run equilibrium relationship between the variables.

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[10] If $r = 0$ then there is no cointegrating relationship between the variables and if $r = 3$ then the three variables are jointly stationary. The rank should therefore be within the range $1 \le r \le 2$.

[11] Since the variables are in logs, normalising on Y gives the elasticities $\hat{\varepsilon}_{Y,S} = -\hat{\beta}_S / \hat{\beta}_Y$ and $\hat{\varepsilon}_{Y,I} = -\hat{\beta}_I / \hat{\beta}_Y$, whilst normalising on S gives, $\hat{\varepsilon}_{S,Y} = -\hat{\beta}_Y / \hat{\beta}_S$ and $\hat{\varepsilon}_{S,I} = -\hat{\beta}_I / \hat{\beta}_S$, and on I gives $\hat{\varepsilon}_{I,Y} = -\hat{\beta}_Y / \hat{\beta}_I$ and $\hat{\varepsilon}_{I,S} = -\hat{\beta}_S / \hat{\beta}_I$.

The question becomes, which is the appropriate variable, $\{Y_t, S_t, I_t\}$ to be used to normalise the vector? All three possible cases are considered and the estimated long run elasticities are reported in Table 5. Since they all have the same maximised log-likelihood value of 109.92 (subject to the single exactly identifying restriction) the size, sign and significance of the estimates will be used to select only one relationship.

Table 5

Estimated Long Run Elasticities

Dependent Variable [1]	**Explanatory Variables** [2]		
	Y	*S*	*I*
Y	–	0.547***	0.154
		(0.212)	(0.249)
S	1.827***	–	–0.282
	(0.707)		(0.561)
I	6.478	–3.547	–
	(10.459)	(7.059)	

Notes: [1] The cointegrating vector was identified by normalising the explanatory variable as the dependent variable.
[2] Standard errors are shown in parentheses and tests of significance are reported assuming normality:
*** Significant at the 1 percent level: ** 5 percent level: * 10 percent level.

The long run relationships between saving and output in the first and second equations are striking. Consistent with the Solow model of economic growth, there is a unique equilibrium relationship between the level of saving and output. The first equation shows a one percent increase in saving is consistent with a 0.55 percent increase in output in long run equilibrium. This estimate is significant at the one percent level (under the assumption of normality).

Consider this in terms of the elasticity of output with respect to the average saving rate (*aps*), $s = S - Y$ (with the variables defined in logs:

$$\varepsilon_{Y,s} = \frac{\Delta Y}{\Delta s} = \frac{\eta}{1-\eta}$$

where η is the share of income that is paid to capital. According to Romer (2006; pp. 22-24) the average share of income paid to capital is around one-third $(\eta \simeq 0.33)$ for most countries, consistent with an elasticity of one-half $\left(\varepsilon_{Y,s} \simeq 0.50\right)$. The estimate of the elasticity in the first normalisation of Table 5 can be modified to incorporate the saving rate, s:

$$Y = 0.55S = 0.55\left(s + Y\right) = \frac{0.55}{1-0.55}s\,.$$

The elasticity estimate of $\varepsilon_{Y,s} = \frac{0.55}{1-0.55} = 1.22$ shows that real GDP has an elastic response to changes in the saving rate on the long run balanced growth path in Iran. This is larger than the inelastic value of 0.50 for most countries. Similarly, the estimate of the share of income paid to capital, $\eta = 0.55$ is also higher than 0.33 for other countries.

The second possible normalisation with saving as the dependent variable in Table 5 gives the inverse elasticity of 1.82.[12] Whilst the direction of the effect of output influencing saving supports the Carroll-Weil hypothesis, the elastic value is large. Inspection of Table 5 clearly shows that investment has no significant long relationship with output and saving. Indeed the determination of investment in the third identified vector is very imprecise, reflecting the variability of investment relative to saving. These results lend strong support for the selection of a rank of one for the system, reflecting the singular, close relationship between output and saving. The first normalisation of output in the first row of Table 5 is selected as the best representation of the long run equilibrium relationship.

The associated short run error correction is therefore:

$$\Delta X_t = -\alpha_X\left(Y_{t-1} - 0.547S_{t-1} - 0.154I_{t-1}\right) + \kappa_X + \delta_X D_t + \sum_{X\in(Y,S,I)} \gamma_X \Delta X_{t-1} + v_{X,t} \qquad (8)$$

where $X \in \{Y, S, I\}$. The results of the estimation of the VECM are summarised in Table 6 and we will focus on the short run growth in output, ΔY_t. The estimated error correction coefficient $\left(\alpha_Y\right)$ has the correct sign and is significant at the one percent level. The magnitude of 0.228 reflects the inertia inherent in the evolution of annualoutput, with nearly 25 percent of disequilibrium eliminated in the first year. Importantly, the inclusion of the saving variable (with significant coefficient at the one percent level) in the error correction means that saving has a short run equilibrating effect on output. The size of the short run elasticity is $-0.125 = 0.228\times(-0.547)$. This further supports the Solow model whereby changes to saving have only transitory effects on the growth in output.

Table 6

Short Run Error Correction Elasticities of Explanatory Variables [1]

	$\alpha_X\ (ecm)$	κ_X	δ_X	$\gamma_Y\ (\Delta Y_{t-1})$	$\gamma_S\ (\Delta S_{t-1})$	$\gamma_I\ (\Delta I_{t-1})$
ΔY_t	0.228 (0.072)***	–0.894 (0.298)***	–0.102 (0.031)***	–0.022 (0.238)	–0.025 (0.044)	0.174 (0.071)**
	$R^2 = 0.50$ $F_{5,36} = 7.32$***		$DW = 2.03$ $F_{1,35}(\rho = 0) = 0.04$		$F_{1,35}(\text{RESET}) = 0.04$ $F_{1,40}(\bar{\sigma}^2) = 2.37$	
ΔS_t	–1.289 (0.309)***	–5.237 (1.274)***	–0.489 (0.131)***	–0.639 (1.021)	–0.060 (0.187)	0.376 (0.304)
	$R^2 = 0.42$ $F_{5,36} = 5.28$***		$DW = 2.23$ $F_{1,35}(\rho = 0) = 4.42$**		$F_{1,35}(\text{RESET}) = 0.02$ $F_{1,40}(\bar{\sigma}^2) = 2.06$	
ΔI_t	–0.211 (0.209)	–0.849 (0.862)	–0.089 (0.089)	0.624 (0.691)	0.054 (0.127)	0.129 (0.206)
	$R^2 = 0.26$ $F_{5,36} = 2.47$**		$DW = 1.78$ $F_{1,35}(\rho = 0) = 2.30$		$F_{1,35}(\text{RESET}) = 0.44$ $F_{1,40}(\bar{\sigma}^2) = 0.02$	

Notes: [1] Standard errors are shown in parentheses and tests of significance are reported assuming normality:
*** Significant at the 1 percent level: ** 5 percent level: * 10 percent level.
$F_{1,35}(\rho = 0)$ represents the Lagrange multiplier test for serial correlation:
$F_{1,35}(\text{RESET})$ represents Ramsey's test using the square of the fitted values:
$F_{1,40}(\bar{\sigma}^2)$ represents the test for heteroscedasticity.

The coefficients on the intercept (κ_Y) and dummy variable (δ_Y) are also significant at the one percent level. The dummy variable coefficient of –0.102 implies that the average growth in output (measured as the first difference in logs, ΔY_t) after 1979 was around ten percent per annum lower than for the period prior to this.

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12 This elasticity is simply given by, $1/0.547 = 1.827$, which must also be significant because the ratios of the coefficients to standard errors must be the same, $0.547/0.212 = 1.827/0.707 = 2.58$.

Importantly, the short run Granger causality test of the lagged dependent variables $\gamma_s \Delta S_{t-1}$ and $\gamma_I \Delta I_{t-1}$ on ΔY_t shows the growth in investment increases the growth in output with the elasticity of 0.174, which is significant at the five percent level. The inclusion of the error correction in the test is important because its exclusion would mispecify the relationship and invalidate the test of short run Granger causality. Note that investment is not important in equilibrating output via the error correction mechanism, because the estimate of 0.154 in the normalised cointegrating vector is not significant. If this was significant then it would support the Solow model of growth, which states that increases in capital only lead to transitory growth in output. In contrast, the estimated Granger causing short run dynamic elasticity of 17.4 percent is consistent with the endogenous growth model whereby increases in capital contribute to sustained growth in output.

The summary statistics show the VECM for the growth in output passes the test for serial correlation (with the DW statistic and the Lagrange multiplier test), Ramsey's RESET test for correct functional form, and the test for heteroscedasticity. Fifty percent of the short run growth in output is explained by the first VECM, which is high given the structural changes that Iran has experienced in this period.

IV Conclusions and Policy Implications

This paper attempts to estimate the interdependencies between capital formation, saving and output for Iran which is complicated for two reasons. The first is the theoretical models and conflicting empirical findings of the relative roles of these important aggregates do not provide clear guidance as to the appropriate specifications. Second, Iran's turbulent history makes it difficult to disentangle the complex and changing interrelationships between output, saving and investment for the period of our study, 1960 to 2003. It is important that structural change in the variables is explicitly incorporated into the simultaneous estimation in a non-stationary growth setting.

The methodology adopted follows the work by Pahlavani (2005) on the causes of economic growth in Iran and uses the procedures adopted by Verma and Wilson (2005) and Chaudhri and Wilson (2000). The Zivot and Andrews (1992) method was used to determine that all three variables were non-stationary, $I(1)$. The endogenously determined time of the most significant structural breaks were 1984 for output, 1980 for saving and

1979 for investment. These years coincide with the effect of the Islamic revolution in 1979 and Iran-Iraq war, 1980 to 1988.

The relationships were estimated using Johansen's (1991, 1995) FIML procedure which is appropriate for estimating the effects of non-stationary variables in a simultaneous setting. The cointegrating vector estimates indicate a long run elasticity of output with respect to saving of 0.55. That is, a one percent increase in saving will be associated with a 0.55 percent increase in the long run equilibrium level of output, which describes a Solow style relationship. This also implies the share of income that is paid to capital in the form of saving in Iran is higher at 0.55 than the average for developed countries of around 0.33. These findings show the importance of saving in promoting higher levels of output and income in Iran. However, whilst they explain a higher long run steady state, they do not explain the causes of economic growth. The role of investment was found to be imprecise in the long run.

The results of the estimation of the short run error correction show that saving has a short run equilibrating effect on output with elasticity of −0.125. This further supports the Solow model whereby changes to saving have only transitory effects on the growth in output. The other important result found that investment dynamically Granger causes the growth in output with a short run elasticity of 0.17, which is significant at the five percent level. This estimate is correctly specified because of the inclusion of the error correction term and the result is consistent with the endogenous growth explanation of growth.

Output is found to return to the equilibrium growth path relatively rapidly, with elasticity indicating around 23 percent of disequilibrium is eliminated in the first year. The structural change parameter in the VCEM estimates that the effect on the growth in output fell by around 10 percent after 1979. This validates the explanation in the introduction that economic growth in Iran slowed significantly after the revolution and war periods.

In summary, the explicit modelling and estimation of endogenously determined structural change in the non-stationary and interdependent measures of output, saving and investment have two important policy implications for Iran. First, whilst relatively high domestic saving is found to be an important determinant of economic growth in the short run and long run, there appears to be scope to reduce the reliance on this domestic source (with the possible use of overseas saving). Second, saving should be used to improve the effectiveness of capital accumulation which was found to be important in promoting economic growth in the short run only. The use of saving in the strategic provision of capital, including infrastructure, is essential for the promotion of long run economic growth in the structurally transforming economy of Iran.

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