

**Temporal Causality between Human Capital and Real Income in
Cointegrated VAR Processes:
Empirical Evidence from China, 1960-1999**

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Abstract

This article examines the causal relationship between human capital and real income using data for China from 1960 to 1999. In the long run there is unidirectional Granger causality running from human capital to real income, while in the short run there is unidirectional Granger causality running from real income to human capital.

Key words: China; human capital; income; cointegration; Granger causality

JEL classification: C12; C22; I2

1. Introduction

There is a large body of literature that examines the correlation between economic growth and real income. Barro (1991), Benhabib and Spiegel (1994), and Barro and Lee (1993) among others find that growth and schooling are positively correlated across countries. These studies have been traditionally interpreted as reflecting the impact of schooling on growth. More recently, Bils and Klenow (2000) have questioned this interpretation, arguing that a plausible alternative explanation is that growth drives schooling rather than that schooling drives growth. Conceptually, causality could run in either direction. On the one hand, education, by increasing the human capital stock of individuals, improves their productivity and therefore contributes to growth. On the other hand, economic growth may provide the resources and surplus needed for further investment in human capital. Moreover, from the viewpoint of the individual, schooling involves sacrificing current earnings for higher future earnings, and economic growth, even when skill neutral, increases the

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wage gains from schooling. Foster and Rosenzweig (1996) argue that this is what happened in India in the 1970s where provinces which benefited from the introduction of Green Revolution technologies saw increases in returns to, and enrollments in, schooling.

However, there are few studies which examine the relationship between economic growth and human capital formation within a causal framework. Extant studies which test for causality between human capital and growth within a bivariate framework are De Meulemeester and Rochat (1995), In and Doucouliagos (1997), and Asteriou and Agiomirgianakis (2001). These studies have mixed results. De Meulemeester and Rochat (1995) test for Granger causality between higher education enrollments and economic growth in six countries: Sweden (1910–1986), United Kingdom (1919–1987), Japan (1885–1975), France (1899–1986), Italy (1885–1986), and Australia (1906–1986). De Meulemeester and Rochat (1995) find unidirectional short-run Granger causality running from higher education enrollments to economic growth in Sweden, the United Kingdom, Japan, and France and neutrality between higher education enrollments and economic growth in Australia and Italy. Meanwhile, In and Doucouliagos (1997) and Asteriou and Agiomirgianakis (2001) find bi-directional causality between economic growth and human capital formation using U.S. annual data (1949–1984) and Greek annual data (1960–1994), respectively.

These studies are for developed countries. There are no studies for large developing countries such as China. This article contributes to the existing literature by employing cointegration and error-correction modeling to test the causal relationship between human capital stock and real income using annual data for China from 1960 to 1999. As is well known, since the start of market reforms in the 1970s China has had one of the highest growth rates in the world. At the same time, while tertiary enrollments are still low relative to other Asian developing countries, the accumulation of basic human capital, at the primary and secondary school level, in China since the Cultural Revolution has also been rapid. The estimated enrollment ratio in primary schools was 25 percent in 1949. This figure increased to 84.3 percent in 1980 and 100 percent in 1996. In 1949 the estimated enrollment ratio in secondary schools was only 2 percent, but this figure had increased to 46 percent by 1977 and was 70 percent in 1996 (Wang and Yao, 2003). There has also been a big commensurate drop in the rate of adult illiteracy. The rate of adult illiteracy has declined from 52 percent in 1964 to 17 percent in 1999 (World Bank, 2001).

The outline of this article is as follows. The next section discusses the data and sets out the econometric methodology and results in three stages: unit root testing, cointegration, and Granger causality tests. The final section concludes the article.

2. The Model

2.1 Data

The data on real income (Y_t) for 1960 to 1995 is extracted from Hsueh and Li

(1999) and updated for the period 1996 to 1999 from the *China Statistical Yearbooks*. The Hsueh and Li (1999) data set is regarded as more reliable than the official estimates of real income up to the 1990s. The human capital index (HC_t) was constructed by Wang and Yao (2003) and is available in an appendix to their article. Wang and Yao (2003) construct a weighted index of educational attainment from five levels of schooling: primary, junior secondary, senior secondary, special secondary, and tertiary. The resulting data set contains new and improved estimates of human capital compared with what was previously available (for full details see Wang and Yao, 2003).

2.2 Unit Root Tests

Prior to estimating Granger causality between real income and human capital, we verified that all the series were integrated of the same order. We first tested for unit roots using the Augmented Dickey Fuller (ADF) and Phillips Perron (PP) tests. The results are reported in Table 1. The ADF and PP statistics for the levels of real income and human capital do not exceed the critical values (in absolute terms). However, when we take the first difference of each of the variables, the ADF and PP statistics are higher than the respective critical values (in absolute terms). Therefore, the ADF and PP tests suggest that the variables are integrated of order one or I(1).

Table 1. Results of ADF and PP Unit Root Tests

Variables	ADF Statistic [LL]	CV	PP Statistic [BW]	CV
$\ln Y_t$	-2.3449 [0]	-3.5297	2.4279 [13]	-2.9389
$\Delta \ln Y_t$	-7.1037 [0]	-2.9411	-15.9656 [37]	-2.9411
$\ln HC_t$	-1.3423 [7]	-3.5577	-0.4467 [4]	-3.5297
$\Delta \ln HC_t$	-3.4019 [8]	-3.2217*	-1.8172 [6]	-1.6115*

Notes: LL is lag length, CV denotes critical values at the 5 percent level, * indicates critical values at the 10 percent level, and BW is the bandwidth. The lag length for the ADF test is selected using the general-to-specific approach proposed by Hall (1991) with a maximum lag length set equal to 8. The bandwidth for the PP test is selected with the Newey-West suggestion using the Bartlett kernel.

Perron (1989) showed that the ADF test has low power when the true data generating process is stationary about a broken linear trend. Wary of the fact that the unit root properties of the human capital and real income series may be influenced by structural changes in the economy, we apply two versions of the Zivot and Andrews (1992) test for a unit root in the presence of a structural break in the trend. We used the Zivot and Andrews (1992) Model A (the “crash model”), which allows for a structural break in the intercept of the trend function, and Model C (the “crash-cum-growth” model), which allows for a structural break in the intercept and slope of the trend function.

Model A has the following form:

$$\Delta y_t = \kappa + \phi y_{t-1} + \beta t + \theta_1 DU_t + \sum_{j=1}^k d_j \Delta y_{t-j} + \varepsilon_t . \quad (1)$$

Model C can be represented as follows:

$$\Delta y_t = \kappa + \phi y_{t-1} + \beta t + \theta_1 DU_t + \gamma_1 DT_t + \sum_{j=1}^k d_j \Delta y_{t-j} + \varepsilon_t . \quad (2)$$

Here Δ is the first difference operator, ε_t is a white noise disturbance term with variance σ^2 , and $t = 1, \dots, T$ is an index of time. The Δy_{t-j} terms on the right hand side of equations (1) and (2) allow for serial correlation and ensure that the disturbance term is white noise. DU_t is an indicator dummy variable for a mean shift occurring at time TB and DT_t is the corresponding trend shift variable, where:

$$DU_t = \begin{cases} 1 & \text{if } t > TB \\ 0 & \text{otherwise} \end{cases}$$

and

$$DT_t = \begin{cases} t - TB & \text{if } t > TB \\ 0 & \text{otherwise.} \end{cases}$$

The breakpoint is searched for over the range of the sample $(0.15T, 0.85T)$. The null hypothesis is that $\phi = 0$ in equations (1) and (2), which implies that the series $\{y_t\}$ is an integrated process without a structural break. The alternative hypothesis is that $\phi < 0$, which implies that $\{y_t\}$ is breakpoint stationary. The breakpoint is selected by choosing the value of TB for which the t-statistic for ϕ is minimized.

While Zivot and Andrews (1992) report asymptotic critical values for this test, they warn that the distribution of the test statistic can deviate substantially from this asymptotic distribution. Thus, we calculate "exact" critical values, which are tailored to our sample size following the Zivot and Andrews (1992) methodology. We estimate an ARMA(p, q) model for each Δy_t , with p and q selected according to the Schwartz Bayesian Criterion (SBC). The implied ARMA process is then used as the data generating process for generation of 5000 sample specific series under the null hypothesis of a unit root with no structural breaks. We then follow Zivot and Andrews (1992) in determination of k and obtain a minimum ADF statistic for each of the 5000 series. The critical values are then constructed from this empirical distribution.

The results for Models A and C for human capital and real income, together with the exact critical values for t_α , are reported in Table 2. The Zivot and Andrews (1992) test provides little additional evidence against the unit root hypothesis. For human capital the absolute value of the test statistics from both Models A and C are less than the critical values at conventional levels of significance. For real income, we are unable to reject the unit root null at the 10 percent level (Model C) and at the 1 percent level (Model A). Taken together, these results confirm the findings from

the ADF and PP tests that both human capital and real income series are I(1) processes.

Table 2. Zivot and Andrews (1992) Test for a Unit Root in the Presence of a Structural Break

	Human Capital		Real Income	
	Model A	Model C	Model A	Model C
<i>TB</i>	1991	1991	1972	1977
α	-0.1030 (-3.6577)	-0.0991 (-2.9548)	-0.3751 (-3.6823)	-0.7299 (-7.0004)
θ	0.0208*** (2.7289)	2.3689 (-0.2219)	-0.0998*** (-3.1881)	-0.0481 (0.0218)
γ	-	0.0220 (-0.0005)	-	-2.0058*** (6.3027)
Lag Length	1	1	1	0
"Exact" Critical Values for t_α				
1%	-6.3905	-6.9127	-6.7369	-7.0820
5%	-5.6999	-6.0840	-5.7310	-5.8797
10%	-5.3845	-5.6855	-5.3524	-5.5698

Notes: *** denotes statistical significance at the 1 percent level. The lag length is selected using the general-to-specific approach proposed by Hall (1991) with a maximum lag length set equal to 8.

2.3 Cointegration

To test for cointegration between real income and human capital stock, we first use the Johansen (1988) approach. There are two Johansen cointegration tests. First, the maximum likelihood estimation procedure provides a likelihood ratio test, called a trace test, which evaluates the null hypothesis of at most r cointegrating vectors versus the general null of p cointegrating vectors. The second likelihood ratio test is the maximum eigenvalue test, which evaluates the null hypothesis of r cointegrating vectors against the alternative of $r + 1$ cointegrating vectors. The reported statistics are for the case where the data generating process has no linear trend but allows a constant term to be confined to the cointegrating relations, although the result is robust to alternative specification of the deterministic variables. We use the SBC to determine the lag length. Because we have annual data, we set the maximum number of lags equal to 2 (see Pesaran and Pesaran, 1997). We use adjusted critical values for the Johansen (1988) test developed by Pesaran et al. (2000), which are automatically reported in the latest version of the MICROFIT™ 4.0 statistical package. These critical values are more precise than alternative critical values reported in Johansen and Juselius (1990) and Osterward-Lenum (1992).

The results for both Johansen cointegration tests are presented in Table 3. The various hypotheses to be tested, from no cointegration (i.e., $r = 0$) to a higher number of cointegration vectors, are reported in the first two columns of the table. The eigenvalues associated with the combination of I(1) levels of the Z_t vector are in the third column, with the statistics ordered from highest to lowest. The critical values are reported at the 95 percent and 99 percent levels of significance in the last

two columns of the table. Starting with the null hypothesis of no cointegration (i.e., $r = 0$) among the variables, the trace statistic is 24.37, which exceeds the 95 percent critical value of 19.96. However, the trace statistic when $r = 1$ is less than the 95 percent critical value. Meanwhile, the maximum eigenvalue test statistic (18.36) for the null hypothesis of no cointegration ($r = 0$) exceeds the 95 percent critical value of 15.67, while the statistic when $r = 1$ (9.24) is less than the 95 percent critical value. Thus, the results of the maximum eigenvalue test corroborate the trace test and we conclude that there is one cointegrating vector between real income and human capital.

Table 3. Results of the Johansen (1988) Tests for Cointegration

Null	Alternative	Statistic	95% Critical Value	99% Critical Value
Maximum Eigenvalue Test				
$r = 0$	$r = 1$	18.36	15.67	20.20
$r \leq 1$	$r = 2$	6.00	9.24	12.97
Trace Test				
$r = 0$	$r \geq 1$	24.37	19.96	24.60
$r \leq 1$	$r = 2$	6.00	9.24	12.97

Notes: Reported statistics are for the optimal lag length, selected using the SBC with the maximum lag length set equal to 2. The estimated coefficient of the long-run cointegrating vector is 2.1222 with a t-statistic of 2.3433, implying that human capital positively and significantly (at the 5 percent level) contributes to real income in China.

A limitation of the Johansen (1988) test is that it does not take into account the effect of a structural break on the long-run relationship between real income and human capital. Therefore, we also implement the Gregory and Hansen (1996) test for cointegration which incorporates a structural break into the cointegrating vector. Gregory and Hansen (1996) propose three models of a structural change.

The first model (denoted C) contains a level shift:

$$y_t = \alpha_1 + \alpha_2 D_t^\tau + \beta x_t + \mu_t, \quad t = 1, \dots, n. \quad (3)$$

The second model (denoted C/T) contains a level shift and trend. It takes the following form:

$$y_t = \alpha_1 + \alpha_2 D_t^\tau + \beta_0 t + \beta_1^\tau x_{2t} + \mu_t, \quad t = 1, \dots, n. \quad (4)$$

Here $D_t^\tau = 0$ for $t < \tau$ and $D_t^\tau = 1$ for $t \geq \tau$. The intercept before the level shift is denoted α_1 , while α_2 is the change in intercept due to the level shift.

The third model (denoted C/S) allows for a regime shift. It takes the form:

$$y_t = \alpha_1 + \alpha_2 D_t^\tau + \beta_0 t + \beta_1^\tau x_{2t} + \beta_2^\tau x_{2t} D_t^\tau + \mu_t, \quad t = 1, \dots, n. \quad (5)$$

Here, α_1 and α_2 are as in equations (3) and (4), β_1 denotes the cointegrating slope coefficients before the regime shift, and β_2 denotes the change in the slope coefficient. To test for cointegration between y_t and x_t with structural change,

i.e., the stationarity of μ_t in equations (3) to (5), Gregory and Hansen (1996) propose a suite of tests. These statistics are the commonly used ADF statistic and extensions of the Z_α and Z_t test statistics of Phillips (1987). These statistics are defined as:

$$ADF^* = \inf_{\tau \in T} ADF(\tau) \quad (6)$$

$$Z_\alpha^* = \inf_{\tau \in T} Z_\alpha(\tau) \quad (7)$$

$$Z_t^* = \inf_{\tau \in T} Z_t(\tau). \quad (8)$$

If the breakpoint is unknown a priori, the model is estimated recursively allowing the breakpoint τ to vary such that $|0.15T| \leq \tau \leq |0.85T|$, where T is the sample size. The null hypothesis of no cointegration is investigated by application of the three tests, i.e., equations (6) to (8). Here we are interested in the smallest values for $ADF(\tau)$, $Z_\alpha(\tau)$, and $Z_t(\tau)$ across all possible breakpoints required to reject the null hypothesis.

The results are presented in Table 4. All the test statistics— ADF , Z_α , and Z_t —support the existence of a long-run relationship between real income and human capital. The ADF and Z_α tests suggest that human capital and real income are cointegrated at the 1 percent significance level, while the Z_t test suggests a cointegration relationship at the 5 percent level of significance across all models.

Table 4. Gregory and Hansen (1996) Test for Structural Change in the Cointegration Relationship

	Real Income and Human Capital					
	ADF^*	T_b	Z_t^*	T_b	Z_α^*	T_b
C	-8.5271*** (K = 0)	1973	-8.6386***	1973	-45.4391**	1981
C/T	-9.1641*** (K = 0)	1978	-12.5836***	1980	-47.9702**	1978
C/S	-8.4669*** (K = 0)	1982	-8.5776***	1982	-48.4691**	1965
Critical Values						
Model [Significance Level]	ADF^* , Z_t^*			Z_α^*		
C [1%]	-5.13			-50.07		
C/T [1%]	-5.45			-57.28		
C/S [1%]	-5.47			-57.17		
C [5%]	-4.61			-40.48		
C/T [5%]	-4.99			-47.96		
C/S [5%]	-4.95			-47.04		
C [10%]	-4.34			-36.19		
C/T [10%]	-4.72			-43.22		
C/S [10%]	-4.68			-41.85		

Note: ** and *** denote statistical significance at the 5 percent and 1 percent levels, respectively.

2.4 Granger Causality

As human capital and real income are cointegrated, we augment the Granger causality test with a lagged error-correction term. Engle and Granger (1987) caution that if the series are integrated of order one, VAR estimation in first differences in the presence of cointegration will be misleading. Following Granger (1969), Y_t is said to be “Granger-caused” by HC_t if the information in the past and present values of HC_t helps to improve the forecast of the Y_t variable, i.e., if $MSE(Y_t|\Omega_t) < MSE(Y_t|\Omega_t')$, where MSE is the conditional mean square root of the forecast of Y_t , Ω_t denotes the set of all relevant information up to time t , and Ω_t' excludes the information in the past and present values of Y_t . The Granger causality test involves specifying a bivariate p th order vector error-correction mechanism (VECM) as follows:

$$(1-L) \begin{bmatrix} \ln Y_t \\ \ln HC_t \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} + \sum_{i=1}^p (1-L) \begin{bmatrix} \beta_{11i} \beta_{12i} \\ \beta_{21i} \beta_{22i} \end{bmatrix} \begin{bmatrix} \ln Y_{t-i} \\ \ln HC_{t-i} \end{bmatrix} + \begin{bmatrix} \theta \\ \vartheta \end{bmatrix} [ECT_{t-1}] + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}. \quad (9)$$

Here, in addition to the variables defined above, α_1 and α_2 denote constant drifts, $1-L$ is the lag operator, ECT_{t-1} represents the lagged error-correction term derived from the cointegrating vector, and ε_{1t} and ε_{2t} are serially independent random errors with mean zero and finite covariance matrix. The dependent variable is regressed against past values of itself and other variables. The optimal lag length p is chosen on the basis of the SBC.

The existence of a cointegrating relationship among real income and human capital suggests that there must be Granger causality in at least one direction, but it does not indicate the direction of temporal causality between the variables. Table 5 examines short-run and long-run Granger causality within the error-correction mechanism (ECM). The Wald F-test of the explanatory variables indicates the significance of the short-run causal effects, while the t-statistics on the coefficients of the lagged error-correction term indicate the significance of the long-run causal effects.

The results suggest that in the long run there is unidirectional Granger causality running from the accumulation of human capital to real income, while in the short run there is unidirectional Granger causality running from real income to human capital. One explanation for this finding might lie in cross-sectional findings for many countries, which suggest that human capital and real income are linked through a non-linear pattern (see Kalaitzidakis et al., 2001). The fact that these nonlinearities are found in a cross-sectional setting using average data over five year periods might explain that in the short run and long run one finds different causality results. These in turn are the result of nonlinearities that characterize the steady state.

Table 5. Results of Granger Causality Tests

Dependent Variable	F-Statistics [Probability]		
	$\Delta \ln Y_t$	$\Delta \ln HC_t$	ECT_{t-1} [t-Statistics]
$\Delta \ln Y_t$	-	0.5281 [0.5948]	-0.0278*** [-4.3509]
$\Delta \ln HC_t$	4.6062** [0.0174]	-	-0.0016 [-0.7108]

Notes: ** and *** denote statistical significance at the 5 percent and 1 percent levels, respectively.

3. Conclusion

In this article we have attempted to address three questions: (1) Is real income and human capital cointegrated in China? (2) Is there causation between real income and human capital, and if so what is the direction of causation? (3) What are the policy implications in the event of causation? To answer the above questions we use the Johansen (1988) and Gregory and Hansen (1996) approaches to cointegration coupled with Granger causality F-tests to ascertain the causality relationships.

Our findings support the existence of a long-run relationship between real income and human capital and provide strong support for the hypothesis that schooling is driving growth. The results suggest that human capital accumulation has been important in explaining real income in the long run, while there are feedback effects from real income to human capital formation in the short run. While China has made great strides in increasing the level of primary and secondary school enrollments, tertiary enrollments at 8 percent in 1998 is still low compared with other Asian developing countries and spending on education lags other countries in the region (Wang and Yao, 2003). Having said this, the tertiary enrollment rate has been increasing and the government has set an enrollment rate in tertiary education of 15 percent for 2005 (Dahlman and Aubert, 2001). The policy implications of our findings are clear. The results suggest that further investment in education and policies designed to further increase enrollment rates will be beneficial for promoting economic growth.

We conclude by offering suggestions for future research. One of the limitations of this study is that we only consider the causal relationship between human capital and real income within a bivariate setting. Future research could consider the relationship between human capital, investment, and real income for China and other countries within a multivariate setting. An alternative is to examine the causal relationship among economic growth, exports, and human capital within a multivariate framework. To this point, with the exception of a study by Chuang (2000) for Taiwan, little attention has been given to studying this relationship. This is in spite of the fact that there is a clear conceptual link between the three, given that endogenous growth theory has argued that either human capital or trade is the primary engine of growth. Further studies which incorporate additional variables, such as exports and investment, will help to illuminate the channels through which human capital causes

growth (and vice-versa) and contribute to broader efforts in the literature to tease out the complex relationship between schooling and economic growth.

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