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"An Empirical Investigation around the Finance-Growth Puzzle in China - With a particular focus on causality and efficiency considerations

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Abstract:

The paper explores a coherent perspective for understanding the multifaceted puzzle of China's financial development. Specifically, it tests competing finance-growth nexus hypotheses using Granger causality tests in a VECM framework for China over the period 1980–2002. The empirical results support a complex set of bidirectional causality between the financial development proxies and economic growth variables. Additionally, bidirectional causality shows the Chinese financial system to be more driven by and closely aligned with real sector activities than exposed to speculative finance. Study findings have several policy implications. Notably, the development of financial institutions should not be emphasized unilaterally. Rather, attention should be given to the complementary and coordinated development of financial reforms and changes in other areas.

Key Words: Financial Development, China, Economic Growth, Financial System Efficiency,

Note: The paper should be considered in progress and is made public to stimulate further discussion and critical comments. Readers' comments are highly appreciated and should be addressed directly to the author.

1. Introduction

A wide body of empirical evidence supports the argument that a well-developed financial system has a positive impact on economic performance by enhancing intermediation efficiency through reduced information, transaction, and monitoring costs. Thus, efficient financial intermediation influences the allocation of resources and productivity growth. Additionally, financial development may enhance economic growth via either capital accumulation or technological changes. The above link suggests that economic growth rarely (if ever) occurs without a well-functioning financial system (see McKinnon, 1973; Shaw, 1973; King & Levine, 1993; Levine et al., 2000; Beck et al., 2000). In other words, if the financial system distorts the allocation of funds and financial repression is in place, then financial depth (as defined by Shaw, 1973) will remain deficient and economic growth will not be sustained.

It follows, then, that in terms of causality, efficient and sound (nonrepressive) financial development leads to economic growth. Yet, despite its theoretical validity, such a view seems inconsistent with recent experience. Specifically, the rapid growth of many Asian economies in the 1970s and 1980s was accomplished despite domestic financial sectors that could not be regarded as developed (Shan et al., 2001), an observation that also holds for China (see Lardy, 1998). With a real GDP growth averaged at 9.4 percent, China's economic performance is extremely difficult to reconcile with the widespread view that its repressive financial system (in the McKinnon-Shaw sense) grossly distorts the optimal allocation of loanable funds and is therefore inefficient. Moreover, China's huge savings rate flies in the face of the conventional wisdom that financial repression artificially creates excess in the demand for credit while discouraging saving.

Whereas the finance-led growth hypothesis is not ruled out in recent studies on the finance-growth link in China (e.g,. Shan & Jianhong, 2006), the literature throws no light on how a repressive and inefficient financial system can successfully generate economic growth. Yet out of this hypothesis emerge three highly relevant questions: How can the apparent paradox be interpreted in light of the finance-growth nexus? Is financial development a prerequisite to economic growth in China or the reverse? Why and how did China's supposedly inefficient financial system accommodate such rapid economic growth?

Given the importance of China in the world economy, and especially the potential implications for the literature on competing financial development strategies, the objectives of this paper are twofold. First, this paper sheds light on the causality in this case by empirically examining the interactions between Chinese financial development and economic growth. Second, the paper attempts to theoretically reconcile the apparent finance-growth puzzle in China by developing some ideas that can provide intuition as to the channel though which financial development and economic growth interact in the Chinese context. The positive correlation between financial development and economic growth is already a stylized fact verified in many studies. However, extant findings on the causal relations and the contribution of the financial sector to (future) economic growth are divergent.

In general, theoretical and empirical studies suggest three types of causal direction between finance and growth. First, in light of China's high savings rate (averaging 38 percent of GDP over the period 1980–93 compared with national investment levels of 37 percent of GDP), the Harrod-Domar growth model[1] would lead to a hypothesis of one-way causality from financial development to economic growth. Any empirical evidence for this hypothesis would imply that China's financial intermediation efficiently allocates resources and sustains higher economic growth. Second, because China's recently begun financial reform followed at least 15 years of strong economic growth resulting from reforms in other areas like the trade sector and state-owned enterprises (Shan, 2003), there should be evidence of unidirectional causality from growth to finance. Such a finding would confirm Shan et al.'s (2001) conclusion that economic growth causes China's financial development. Nonetheless, a third alternative, the coevolution (bidirectional causality) between economic growth and financial development hypothesized in both early and recent literature (Gurley & Shaw 1960, 1967; Bencivenga & Smith, 1998), cannot be ruled out.

Empirical determination of the relevant causal direction usually resorts to the standard Granger noncausality test (1969); however, results from these tests are highly sensitive to the order of lags in the autoregressive process. That is, choosing an inadequate lag length leads to inconsistent model estimates, and any inferences are likely to be misleading. Moreover, from an economic viewpoint, there is no compelling theoretical support for the lag lengths for all variables in all equations being symmetric.

To address the above concern, this present analysis resorts to Hsiao's (1981) version of the Granger noncausality tests, which uses a cointegration and error correction framework. By avoiding arbitrary lag length selection, as well as the use of symmetric lags in conventional VAR models, Hsiao's approach ensures more reliable results than those in many previous studies on the finance-growth nexus. To the best of our knowledge, no finance-growth nexus study addressing Granger causality shortfalls exists at present. Therefore, in this sense, this paper represents an advance in the current debate.

The remainder of the paper is organized as follows. Section 2 briefly reviews selected theoretical and empirical literature, after which Section 3 discusses methodological considerations and the data. Section 4 empirically tests the competing hypotheses of whether finance causes growth or growth causes finance in China. Section 5 discusses the relevance of the findings, and Section 6 draws out the policy implications for China's financial development strategy.

[1] This model assumes that the growth of output in the current year is proportional to the investment ratio (the share of investment in output) in the previous year.

2. Literature on economic growth and financial intermediation

A general consensus exists among economists that a well-functioning financial sector spurs economic growth (Schumpeter, 1911; Levine, 1997). Major theoretical literature on financial development and economic growth processes postulate four distinguishable, but not mutually exclusive, effects of financial activity and development on overall economic performance: The first is the provision of an inexpensive and reliable *means of payment*; the second, a *volume and allocation effect*, in which financial activity increases resources that can be channeled into investment while improving the allocation of resources devoted to investment. The third is a *risk management effect* by which the financial system helps to diversify liquidity risks; thereby enabling the financing of riskier but more productive investments and innovations (Greenwood & Jovanovic, 1990; Bencivenga & Smith, 1991). The fourth is an *informational effect*, according to which ex ante information about possible investment and capital are made available; ameliorating—although not necessarily eliminating—the effects of asymmetric information (Levine, 2004).

From an aggregate production function point of view, each of these financial functions may contribute to the transformation of a given amount of savings and investment inputs into a larger amount of output through either a capital accumulation channel (Hicksian type; Hicks, 1969) or a technological change channel (Schumpeterian type; Schumpeter, 1911). Despite Schumpeter's (1911) argument that well-functioning financial institutions spur technological innovation, early growth models (including the Solow-Swan) did not incorporate the role of financial systems because economic growth theorists believed technological progress and population growth to be the main driving forces behind economic growth and did not see financial systems as directly relevant. Rather, in neoclassical exogenous growth theory, financial intermediation was thought to influence growth only via saving channels. Thinking changed with the development of endogenous growth models in which financial intermediation plays a more specific role through the financing of R&D, as well as investment in human capital (see e.g., Barro & Sala-I-Martin, 1995; Romer, 1986), and the existence of externalities. Such endogenous growth models allow finance to function as a cause of technological progress and capital accumulation, which can in turn accelerate economic growth.

Early theoretical considerations on financial system development (Gurley and Shaw, 1955; Patrick, 1966; and Goldsmith, 1969) show divergent patterns in the link between financial sector and growth. For instance, in the two-way relationship between financial development and economic growth postulated by Lewis (1955), financial markets develop as a consequence of economic growth and then act as a stimulant to real growth. This view is supported by Patrick (1966) who hypothesizes two alternatives of finance-growth interactions. First, the development of financial markets promotes economic growth (the supply leading approach) by reducing market imperfections and frictions. Simultaneously, in a reverse channel from economic growth to financial development, economic growth produces an increased demand for financial services, meaning that financial development must be understood as a demand driven phenomenon (the demand driven approach).

Somewhat earlier, Robinson (1952) argued that where enterprise leads, finance follows. That is, rising income levels create demands for particular types of financial arrangements from households and business sectors, and the financial system responds automatically to these demands. Additionally, Goldsmith (1969) maintains that the process of growth has feedback effects on financial markets by creating incentives for further financial development. The two-way relationship between financial development and economic growth is supported by a number of endogenous growth models (Greenwood & Jovanovic, 1990; Berthelemy & Varoudakis, 1997; Greenwood & Bruce, 1997).

Also receiving considerable empirical support in contemporary studies is the positive association between financial systems and economic growth in which the level of financial development is a good predictor of economic growth (see especially, King & Levine, 1993; Rousseau & Wachtel, 1998). Most of these studies conclude that higher levels of financial development are significantly and robustly correlated with faster current and future rates of economic growth, physical capital accumulation, and economic technological change (see Bencivenga & Smith, 1991; Bencivenga et al., 1995; Greenwood & Jovanovic, 1990, among others). However, Demetriades and Hussein (1996), De Gregorio and Guidotti (1995), and Odedokun (1996) moderate this claim by emphasizing that such effects differ across countries, time periods, and/or stages of development. Moreover, Gregorio and Guidotti (1995) specify that the mechanism of financial development on economic growth is due to its impact on “efficiency” rather than the “volume” of investment.

Beyond the evidence that the level of financial development is a good predictor of future rates of economic growth (Levine, 1997; King & Levine, 1993), Patrick's (1966) problem—that is, which is the cause and which the effect—remains unsolved (McKinnon, 1988). Is finance a leading sector in economic development, or does it simply follow growth in real output generated elsewhere?

Unfortunately, there is no simple procedure to determine which view is empirically adequate – not even one that would rule out some views as obviously false (Graff, 2001). For instance, Rousseau and Wachtel (1998) find one-way causality between financial development and economic growth in the case of five OECD countries during an earlier period of fast industrialization (1871-1929), while a panel data analysis by Beck et al. (2000) shows that banks have a strong causal effect on economic growth. In contrast, King and Levine (1993b) conclude that the initial level of financial development predicts future growth rate. However, using time series analysis, Arestis and Demetriades (1997) conclude that the evidence favors a bidirectional relationship between financial development and economic growth. Moreover, Murende and Eng (1994), Demetriades and Hussein (1996), and Luintel and Khan (1999) find evidence of bidirectional causality between financial development and economic growth in all their sample countries.

The above theoretical and empirical evidence suggests that economic growth occurs within a well-functioning financial system (see McKinnon, 1973; Shaw, 1973). In other words, if the financial system distorts the allocation of funds and financial repression is in place, then financial depth (as defined by Shaw, 1973) will remain deficient and economic growth will not be sustained. It follows that removing the government intervention and leaving price mechanism to the market are the key preconditions of economic growth.

The Shaw-McKinnon argument rests on the view that markets allocate resources best and that state development management agencies are not better at plotting development paths. The argument ignores some of the structural features commonly found in developing economies. For instance, asymmetric information and externalities in financial markets (Stiglitz & Weiss, 1992) can lead to sub-optimal levels of financing and investment, an inefficient allocation of capital, or have other undesirable consequences such as illiquidity which are detrimental for economic growth. Some of these market imperfections may be best addressed through appropriate oversight by the government. The degree of government involvement in regulating and otherwise influencing the financial system determine the degree to which the latter will make a positive contribution to growth (Lawrence, 2003). In this regard, analyzing the case of China, Cull and Xu (2000) found that the link between bank finance and subsequent productivity suggest that banks were somehow able to identify and lend to relatively productive state owned enterprises.

On the issue of causality between finance and economic growth there have been some disturbing empirical results. For instance, an earlier study by Aziz and Duenwald (2002) concludes that the positive link between finance and growth in China is more apparent than real in that the nonstate sector, which contributed most of China’s remarkable growth, did not resort to the domestic financial system in any substantial way for financing. Even more disturbing results are provided by Boyreau-Debray’s (2003) study on Chinese financial intermediation and growth, which finds that credit extended by the banking sector at the state level has a negative impact on provincial economic growth. In the same vein, De Gregorio and Guidotti (1995) find evidence for a negative relationship between financial development and growth in 12 Latin American countries during the period from 1950 to 1985.

Overall, empirical studies on Chinese finance-growth dynamics are not only limited in number but have produced no consensual interpretation of efficiency at a macroeconomic level. Such studies usually measure financial system efficiency in terms of allocative efficiency, meaning it can be judged either directly by monitoring some proxy of allocative efficiency or indirectly by estimating the contribution of a financial variable to economic growth. Allocative efficiency can also be inferred indirectly by studying whether a bank's resources are allocated to the most productive uses or not. Most productive use, in turn, can be defined in terms of the macroeconomic rate of return proxied by GDP growth rate. Thus, the causal chain between economic growth and financial development in China remains ambiguous and, together with the debate on China's macrofinancial efficiency, merits an alternative investigation using a nonstandard methodological approach.

3. Model, Methodological Considerations, and Data

The section presents the empirical framework and discusses the unit root and cointegration test procedures, causality tests, and datasets.

3.1. Standard Empirical Framework

Drawing from Pagano (1993) and Bailliu (2000), the relationship between capital flows and growth can be examined using a simple endogenous-growth *AK* model. The potential effects of changes in financial variables (i.e., financial development) on steady-state growth through their influence on capital accumulation, in a closed-economy, is given by:

$$Y_t = AK_t \tag{1}$$

where output is a linear function of the aggregate capital stock. This type of production function can be seen as a reduced form for either a framework in which the economy is competitive with external economies, as in Romer (1989), or one in which *Kt* is assumed to be a composite of physical and human capital, as in Lucas (1988), where the two types of capital are reproducible with identical technologies. There is no population growth in this model and the economy produces only one good, which can be consumed or invested. By assuming that the capital stock depreciates, gross investment equals:

$$I_t = K_{t+1} - (1-\delta)K_t \tag{2}$$

Financial intermediation consists in the process of transforming savings into investment. The transaction cost involved $(1-\delta)$ can be seen as the spread between lending and borrowing rates charged by banks. Capital market equilibrium requires that available savings (gross savings minus transaction costs) be equal to gross investment. Thus, equilibrium in the capital market ensures that

$$\phi S_t = I_t \tag{3}$$

Using equations (1) through (3) and dropping the time indices, the growth rate of output, *g*, can be written as follows

$$g = A\left(\frac{I}{Y}\right) - \delta = A\phi s - \delta \tag{4}$$

where *s* denotes the gross savings rate. Equation (4) thus represents the steady-state growth rate of the model with financial intermediation and reveals two main channels through which financial development can affect economic growth. Financial development is assumed to occur as a result of increased financial intermediation, although it could also be influenced by other factors—such as financial innovation or government policies. The first channel involves the efficiency with which savings are allocated to investment. As banks engage in increased intermediation, they are likely to become more efficient at what they do, and thus the spread between their lending and borrowing rates falls. This result in an increase in the proportion of savings channeled to investments; thus, g in equation (4) will increase as a result of an increase in ϕ. Second, an increase in financial intermediation can affect growth if it improves the allocation of capital; which in this model is translated into higher growth, because it increases the overall productivity of capital, A.

Based on the above, we now set out a simple model to test the hypothesis that financial development is linked to economic growth. The simplest relevant growth model is the *AK* production function in which aggregate output is a linear function of the aggregate physical capital stock. Hence, the finance-growth relationship can be represented as commonly found in the literature (e.g., Demirguc-Kunt & Levine, 2001):

$$Y_t = \alpha + \beta X_t + \gamma Z_t + \mu_t \tag{5}$$

where Y_t is the growth of per capita GDP for some time period, t; X_t indicates a set of measures of financial sector development; and Z_t represents a conditioning variable. Achieving the study goal requires a three-phase process: an analysis of the integration order of the variables; a test for cointegration among time series; and implementation of Hsiao's version (1981) of the Granger noncausality method (Granger, 1969) to estimate causality for each equation of the model.

3.2. Unit root and cointegration testing procedure

Nonstationary time series Y_t is said to be integrated of order d, [$Y_t \sim I(d)$], if it achieves stationarity after being differenced d times (Granger, 1986; Engle & Granger, 1987). To determine the order of integration, the most common unit root test is the Dickey-Fuller (DF) or augmented Dickey-Fuller test (ADF; Dickey & Fuller, 1979, 1981), which estimates the following equation:

$$\Delta y_t = c_1 + \varpi y_{t-1} + c_2 t + \sum_{i=1}^{\rho} d_i \Delta y_{t-i} + v_t \tag{6}$$

In (6), $\{y_t\}$ is the relevant time series, Δ is a first-difference operator, t is a linear trend, and v_t is the error term. The above equation can also be estimated without including a trend term (by deleting the term $c_2 t$). The null hypothesis of the existence of a unit root is H_0: $\omega = 0$.

Once the time series is ascertained to be integrated of the same order, for example, I(1), it should be examined for cointegration. Cointegration regressions measure the long-term relationships between the variables whose existence guarantees that the variables demonstrate no inherent tendency to drift apart. We use the Johansen cointegration tests (Johansen 1988; Johansen & Juselius, 1990), which set up the nonstationary time series as a vector autoregression (VAR) of order p:

$$\Delta y_t = \mathbf{\Pi} y_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta y_{t-i} + \mathbf{B} x_t + \varepsilon_t \tag{7}$$

$$\mathbf{\Pi} = \sum_{i=1}^{p} A_i - I, \quad \mathbf{\Gamma} = -\sum_{j=i+1}^{p} A_j. \tag{8}$$

where y_t is a k-vector of the I(1) variables, x_t is a vector of the deterministic variables, and ε_t is an identically and independently distributed error term. The rank of the coefficient matrix, $\mathbf{\Pi}$, is reduced if $r < k$, where r is the number of cointegrating relations. In this case, there exists $k \times r$ matrices $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$, each with rank r such that $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$ and $\boldsymbol{\beta}' y_t$ is stationary. The matrix $\boldsymbol{\beta}$ is the matrix of cointegrating parameters, and the matrix $\boldsymbol{\alpha}$ is the matrix of weights with which each cointegrating vector enters the k equations of the VAR.

Johansen and Juselius (1990) provide two different statistics to test the hypothesized existence of r cointegrating vectors: the trace test statistic and the maximum eigenvalue test (Johansen, 1988; Johansen & Juselius, 1990). The trace test statistic tests the null hypothesis that the number of distinct cointegrating vectors is less than or equal to r against a general alternative. Alternatively, the likelihood ratio (LR) statistic, known as the maximum eigenvalue test statistic, tests the null hypothesis that the number of cointegrating vectors is r against the alternative of $r+1$ cointegrating vectors. Thus, cointegration is a necessary step in determining the most appropriate specification for the causality test.

3.3. Causality test procedure

According to Granger (1969), if the inclusion of past (lagged) values of X significantly contributes to the explanation of Y in a regression of Y on its own past values and all other relevant information, then X is said to Granger cause Y. To examine the nature of the causality between the Y and X series, an appropriate Granger causality test requires determination of an equal lag length VAR involving Y and X. An inadequate choice of the lag length would produce inconsistent model estimates, and any inferences would probably be misleading. The importance of lag length determination is demonstrated by Braun and Mittnik (1993), who show that estimates of a VAR whose lag length differs from the true lag length are inconsistent. Lutkepohl (1993) also demonstrates that overfitting (selecting a higher order lag length than the true lag length) causes an increase in the VAR mean square forecast errors, whereas underfitting the lag length often generates autocorrelated errors.

Like most VAR models, the Granger noncausality tests are estimated using symmetric lags (i.e., the same lag length is used for all variables in all equations of the model); however, as previously mentioned, economic theory provides no compelling reason that lag lengths must be symmetric. Thus, in response to concerns about arbitrary lag determination and symmetric lags, Hsaio (1981) suggests estimating VARs in which the lag length on each variable in each equation can differ. Hsiao's approach also combines the Granger concept of causality and Akaike's final prediction error criterion (Akaike, 1969), and is specifically designed to avoid the imposition of false or spurious restrictions on the model. For a detailed discussion of Hsiao's version of the Granger causality method, see Hsiao (1981, 1982), Cheng and Lai (1997), and Bajo-Montavez (2002).

Hsiao's variant of the Granger causality test can best be illustrated by a practical example. Assuming that the two stationary variables *Yt* and *Xt* must be tested for Granger causality, we consider two models

$$\mathrm{Y_t} = \alpha + \sum_{i=1}^{m} \beta_i \mathrm{Y_{t-i}} + \mathrm{u_t} \tag{9}$$

$$\mathrm{Y_t} = \alpha + \sum_{i=1}^{m} \beta_i \mathrm{Y_{t-i}} + \sum_{j=1}^{n} \gamma_j \mathrm{X_{t-j}} + \mathrm{v_t} \tag{10}$$

where α is a constant term, β and γ are coefficients of exogenous variables, and u_t and v_t are white noise error terms with the usual statistical properties. Hsiao's procedure then involves the following steps:

(i) Y_t is assumed to be a univariate autoregressive process as in (9), and its final prediction error criterion (FPE) is computed with the order of lags *i* varying from 1 to *m*. The lag *m* that that yields the smallest FPE is selected, and its corresponding FPE is denoted as $\mathrm{FPE}_Y(m, 0)$.

The corresponding FPE is given by

$$FPE(m) = \frac{(T+m+1)}{T-m-1} x \frac{SSE}{T} \tag{11}$$

where *T* denotes the number of observations in the regression, and *SSE* is the sum of squared residuals. Causality can then be determined as follows.

(ii) *Yt* is treated as a controlled variable with *m* lags, then the lags of *Xt* are added to (9) as in (10), and the FPEs are computed with the order of lags *j* varying from 1 to *n*. The lag *n* that yields the smallest FPE is selected, and its corresponding FPE is denoted as $\mathrm{FPE}X(m, n)$.

The corresponding FPE is given by

$$FPE(m^*, n) = \frac{(T+m^*+n+1)}{T-m^*-n-1} x \frac{SSE(m^*, n)}{T} \tag{12}$$

(iii) $\mathrm{FPE}_Y(m, 0)$ is then compared with $\mathrm{FPE}_Y(m, n)$. If $\mathrm{FPE}_Y(m,0) > \mathrm{FPE}_Y(m, n)$, then X_t is said to Granger-cause Y_t, whereas if $\mathrm{FPE}_Y(m, 0) < \mathrm{FPE}_Y(m, n)$, then Y_t is not Granger-caused by X_t.

Reverse causality (whether Y_t Granger causes X_t) is determined by repeating steps (i) to (iii) with X_t as the dependent variable.

In practice, the implicit assumption that Y_t and X_t are stationary must be confirmed before (9) and (11) can be implemented. If the series are nonstationary with unit roots, they must be transformed into stationary ones by means of a difference filter. If the variables are all integrated of the same order—for example, I(1)—a check should be run for cointegration.

Such cointegration would imply that any standard Granger causal inferences will be invalid unless an error correction mechanism (ECM) is included.

Engle and Granger (1987) demonstrate that once a number of variables (e.g., *Y* and *X*) are found to be cointegrated, there always exists a corresponding error correction representation, which implies that changes in the dependent variable are a function of the level of disequilibrium in the cointegration relationship (captured by the error correction term) as well as of changes in other explanatory variable(s). A consequence of ECM is that either ΔYt or ΔXt or both must be caused by the value of the previous period error term derived from the cointegrating equation. Intuitively, if *Y* and *X* have a common trend, then the current change in *Y* (e.g., the dependent variable) is partly the result of *Y* moving into alignment with the trend value of *X* (e.g., the independent variable). Through the error correction term, the ECM opens up an additional channel (ignored by the standard Granger tests) through which Granger causality can emerge. Consequently, (9) and (10) should be modified to incorporate an error correction mechanism, derived as follows from the residuals of the appropriate cointegration relationship:

$$\Delta \mathrm{Y}_{\mathrm{t}} = \alpha + \sum_{i=1}^{m} \beta_i \Delta \mathrm{Y}_{\mathrm{t-i}} + \delta z_{t-1} + \mathrm{u}_{\mathrm{t}} \tag{13}$$

$$\Delta \mathrm{Y}_{\mathrm{t}} = \alpha + \sum_{i=1}^{m} \beta_i \Delta \mathrm{Y}_{\mathrm{t-i}} + \sum_{j=1}^{n} \gamma_j \Delta \mathrm{X}_{\mathrm{t-j}} + \delta \mathrm{z}_{\mathrm{t-1}} + \mathrm{v}_{\mathrm{t}} \tag{14}$$

where z_{t-1} is the vector error correction term (Engle & Granger, 1987), which stands for the short-term adjustment to long-run equilibrium trends. It should be noted that if X_t and Y_t are I(1) but not cointegrated, no error correction mechanism binds the two variables and there is no one-period lagged error term in (13) and (14).

The introduction of a control variable, however, demands some modification of the VAR equation. Specifically, testing for Granger causality in the trivariate case requires that (13) and (14) be amended by the adding of a third variable, *W*, to give the following model

$$\Delta \mathrm{Y}_{\mathrm{t}} = \alpha + \sum_{i=1}^{m} \beta_i \Delta \mathrm{Y}_{\mathrm{t-i}} + \sum_{k=1}^{p} \theta_{\mathrm{k}} \Delta W_{t-k} + \delta z_{t-1} + \mathrm{u}_{\mathrm{t}} \tag{15}$$

$$\Delta \mathrm{Y}_{\mathrm{t}} = \alpha + \sum_{i=1}^{m} \beta_i \Delta \mathrm{Y}_{\mathrm{t-i}} + \sum_{j=1}^{n} \gamma_j \Delta \mathrm{X}_{\mathrm{t-j}} + \sum_{k=1}^{p} \theta_{\mathrm{k}} \Delta W_{t-k} + \delta z_{t-1} + \mathrm{v}_{\mathrm{t}} \tag{16}$$

with its corresponding FPE:

$$FPE(m^*, n, p) = \frac{(T + m^* + n^* + p + 1)}{T - m^* - n^* - p - 1} \, x \, \frac{SSE(m^*, n^*, p)}{T} \tag{17}$$

In the trivariate case, the relevant comparison is between $\mathrm{FPE}_{\Delta Y}(m, 0, p)$ and $\mathrm{FPE}_{\Delta Y}(m, n, p)$, where (*m*, 0, *p*) and (*m*, *n*, *p*) are the combinations of lags leading to the smallest FPE in each case. If $FPE_{\Delta Y}(m, 0, n) > FPE_{\Delta Y}(m, n, p)$, X Granger causes Y conditional on the presence of the third variable W.

3.4. Data and stationarity tests

The sources for all series data, which cover the period from 1980 to 2002, are the IMF publication *International Financial Statistics* (CD ROM, 2004) and the World Bank's *World Development Indicators* (2003). Specifically, the data are taken from three indexes of financial development: financial deepening, proxied by liquid liabilities, *L*, (M3/GDP); the credit extended to the private sector by banks (as a percentage of the GDP); and the ratio of total credit extended to the entire economy by the banking sector (also as a percentage of the GDP).

The first indicator *M* (M3/GDP) expresses financial intermediary development and measures the liquid liabilities of the financial system (currency plus demand and the interest bearing liabilities of the bank and nonbank financial intermediaries) divided by GDP. As argued in De Gregorio and Guidotti (1995), monetary aggregates like M3 may be good proxies of financial development because they are highly related to both the ability of financial systems to provide transaction services and the ability of financial intermediaries to channel funds from savers to borrowers. Moreover, because the role of capital markets in China, as in other developing economies, is unusually small, authors such as Gelb (1989) and King and Levine (1993) use M3.

Again following King and Levine (1993), the second and third indicators are the ratio of credit from banks to the private sector as a share of the nominal GDP and the ratio of total credit from banks to the economy as a share of the nominal GDP. The former, *PC,* measures the value of credits from financial intermediaries to the private sector divided by GDP. It excludes credits issued by central and development banks, credit to the public sector, and cross claims of one group of intermediaries on another. The latter, *C*, combines the credit provided by banks to both the public and private sectors. Because the ratio of bank credit to GDP is directly linked to investment and economic growth, the credit provided to the economy is assumed to generate increases in investment and productivity.

Nonetheless, De Gregorio (1996) argues that even though the credit/GDP indicator is a good indicator of financial development occurring through the banking system, it may be a weak indicator of financial development taking place outside the banking system—for example, in the stock markets (De Gregorio & Guidotti, 1995) or through informal or self financing. However, this weakness may be less relevant in countries such as China, in which most financial development occurs within the banking system. Moreover, since total credit was largely dominated by directed lending (80 percent), the indicator could also be interpreted as a proxy for financial restraint policies.

As regards the use of the ratio of total credit to GDP as a proxy for macrofinancial efficiency, it should be noted that in a strict macroeconomic sense, an efficient financial system should be able to channel a greater volume of funds towards productive investment, thereby boosting economic growth. Thus, the focus should be on *macroeconomic* allocative efficiency. In other words, credit to the economy can also be interpreted as a measure (albeit an imperfect one) of macro efficiency.

Following Levine (1997), economic growth is proxied by the logarithm of *Y*, the annual series of per capita GDP growth. Also commonly added into this type of study are variables for controlling the possible effects of other growth-determining factors like measure of openness to trade and external financing variables (Levine, 1997). Therefore, our model includes FDI flows to control for the external factors associated with the magnitude of GDP growth fluctuations in China. Foreign direct investment (*F*) measures the net inflow of investment to acquire a lasting management interest (10 percent or more of the voting stock) in an enterprise operating in China. It is the sum of equity capital, reinvestment of earnings, other long-term capital, and short-term capital as shown in the balance of payments. All variables in the dataset are transformed into natural logarithms so they can be interpreted in growth terms once the first difference is taken.

4. Estimation results

This section outlines the results from the stationarity tests and Hsiao's version of the Granger causality test, respectively.

4.1. Results from the stationarity and cointegration tests

Before the cointegration tests can be performed, it must be established that the variables are integrated processes of the same order. Therefore, all five variables, *lnG*, *lnM*, *lnC*, *lnCP* and *lnF*, are subjected to the Dickey-Fuller and augmented Dickey-Fuller tests (Dickey and Fuller; 1979, 1981). The ADF regression and null hypothesis of a single unit root cannot be rejected at the 10 percent level for any variable, and each of the five series becomes I(0) after first differencing. Table 1 shows the results at the 5 percent level for *lnCT*, *lnM*, and *lnF*; and at the 10 percent level for *lnG* and *lnPC*.

Table 1: Unit Root Test Results				
	LEVEL		**DIFFERENCE**	
	Without time trend			
	DF	ADF	DF	ADF
G	-1.12	-1.45	-1.75**	-2.78*
C	0.69	0.35	-4.06*	-4.01*
M	0.14	-0.25	-4.54**	-3.82**
PC	0.157	-0.461	-4.46**	-4.43**
F	-0.2	0.26	-2.29*	-2.48*
	With time trend			
G	-4.18*	-4.23	-3.60*	-3.18**
C	-2.20	-1.91	-4.24*	-3.98*
M	-2.41	-2.80	-4.43*	-4.32*
PC	-2.66	-2.54	-4.54**	-4.30*
F	-2.09	-1.94	-2.88**	-2.77**

Note: (*) and (**) indicate respectively the level of significance at 5% and10%.

We test for the number of cointegrating vectors using one specification based on the assumption that the series have a linear deterministic trend and the cointegrating equations have intercepts. We determine the required lag length using the Akaike information criterion (AIC), which selects the number of lags required in the cointegration test. After a VAR model is first fitted to the data, the AIC gives lag 2 as the appropriate lag structure for G-M-F cointegration, PC-G-F cointegration, and G-M-F cointegration. We also test for the couples G-F, PC-F, M-F, and C-F using bivariate cointegration.

The max(λ) and the trace statistic (Johansen, 1988; Johansen & Juselius, 1990) for this model are presented in Tables 2 (a-g). The null hypothesis of the absence of a cointegrating relation between the endogenous variables is rejected at the 95 percent confidence level for both statistics. However, the null of the existence of only one cointegrating vector cannot be rejected at the 90 percent level for either statistic. The presence of cointegration between the financial development and economic growth variables confirms the existence of a long-term relationship among the variables and is consistent with the theoretical predictions of finance-growth theories.

TABLE 2(a): Johansen Cointegration Test Results for GDP, Money and FDI (G-M-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	34.68	25.82	r=0	r>1	62.83	42.91
r=1	r=2	16.17	19.38	r≤1	r>2	28.14	25.87
r=2	r=3	11.97	12.51	r≤2	r>3	11.97	12.51

TABLE 2(b): Johansen Cointegration Test Results for Private Credit, GDP and FDI (PC-G-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	30.54	25.82	r=0	r>1	49.51	42.91
r=1	r=2	14.77	19.96	r≤1	r>2	18.96	25.87
r=2	r=3	7.33	12.51	r≤2	r>3	4.18	12.51

TABLE 2(c): Johansen Cointegration Test Results for Total Credit, GDP and FDI (C-G-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	29.13	24.25	r=0	r>1	46.86	35.01
r=1	r=2	16.90	17.14	r≤1	r>2	17.72	18.39
r=2	r=3	0.75	3.84	r≤2	r>3	0.75	3.84

TABLE 2(d): Johansen Cointegration Test Results for GDP and FDI (G-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	17.99	11.22	r=0	r>1	20.06	12.32
r=1	r=2	2.07	4.12	r≤1	r>2	2.07	4.12

TABLE 2(e): Johansen Cointegration Test Results for Private Credit and FDI (PC-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	12.08	11.22	r=0	r>1	14.37	12.32
r=1	r=2	2.28	4.12	$r\leq 1$	r>2	2.28	34.12

TABLE 2(f): Johansen Cointegration Test Results for Money and FDI (M-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	13.63	11.22	r=0	r>1	16.52	12.32
r=1	r=2	2.89	4.12	$r\leq 1$	r>2	2.89	4.12

TABLE 2(g): Johansen Cointegration Test Results for Total Credit and FDI (C-F)

Maximal Eigenvalue Test				Trace Test			
Null H_0	Alternative H_1	λ_{max}	Critical Value (95%)	Null H_0	Alternative H_1	τ_{max}	Critical Value (95%)
r=0	r=1	13.32	11.22	r=0	r>1	16.26	12.32
r=1	r=2	2.94	4.12	$r\leq 1$	r>2	2.12	4.12

We thus estimate the ECM under the assumption of only one cointegrating equation (CE). Because all signs of the estimates of the CE parameters are as expected, signaling the presence of a cointegrating relationship in each set of variables, we can proceed with the causality analysis using equations (15) and (16) to capture information on a long-term relationship between the level variables.

4. 2. Results from Hsiao's version of Granger causality

As explained in Section 3.3, the Hsiao version of the Granger noncausality test allows each variable to enter the VAR with its own lag length. We determine individual series' lag length using Akaike's FPE criterion. Table 3 reports the minimum FPEs for the three univariate autoregressions with *G* at lag 3; *C* , *M*, and *CP* at lag 1; and *F* at lag 2.

Table 3. Final Prediction Error (FPE) of One-Dimensional AR Processes

Order of Lags	FPE of *lnG*	FPE of *lnC*	FPE of *LnM*	lnCP	FPE of *LnF*
0	0.20960	0.0796	0.1153	0.0511	2.2490
1	0.00087	0.0055*	0.0034*	0.0045*	0.0823
2	0.00074	0.0065	0.0037	0.0051	0.0683*
3	0.00058*	0.0066	0.0035	0.0052	0.0689

(*) indicates lag order selected by FPE criterion at 5% level

Taking into account the cointegration evidence, causality is established by comparing the minimum FPE derived from a bivariate (equation 15) and trivariate VAR (equation 16). The results of Hsiao's variation of the Granger test are presented in Table 4 with the error correction term under the null hypothesis of noncausality. As the table shows, in the growth equation, the FDI (*F*) is added as the first manipulated variable (step 1), after which *C* is added to the previous equation (step 2). Since the FPE obtained in the first step is smaller than that obtained in the second step, the hypothesis that total credit (*C)* does not Granger cause economic growth (*G*) can be rejected.

A similar procedure is implemented for total credit equation, private credit equation, and money equation, respectively. The results, outlined in Table 4 and presented graphically in Figure 1, can be summarized as follows: unidirectional causality is identified running from growth to money and from growth to private sector credit; however, bidirectional causality is found between economic growth and total credit. Similar conclusions are drawn for the error correction terms (*ECT*), which the results show to be negative and statistically significant at either the 1 percent or 5 percent level in all instances but the *G-lnPC* equation. The estimated coefficients range from –0.287 (for the *G-C* equation) to –0.847 (for the *PC-G* equation), indicating immediate convergence to long-run equilibrium after a shock. The regressions fit reasonably well and generally pass the diagnostic tests against serial correlation of the first and fourth order, heteroskedasticity, and structural stability.

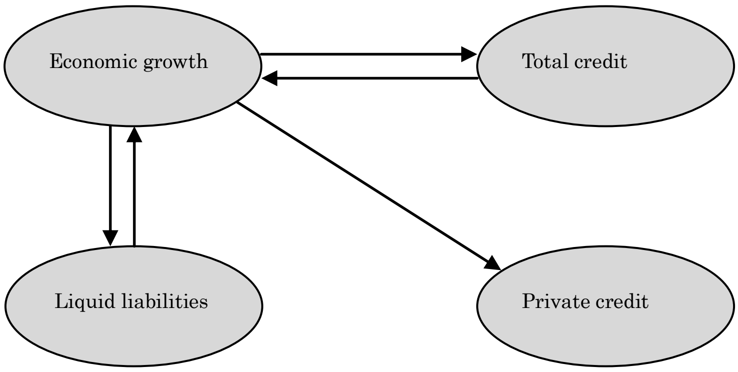
The results shown in Table 4 are further confirmed by the joint *F*-statistics, which reveal that the coefficients in each of the trivariate equations are significant at either the 5 percent or 10 percent level. Our results largely corroborates those by Shan and Jianhong (2006), which not only finds bidirectional causality between financial development and economic growth but also concludes that the Granger causality from economic growth to financial development is stronger than that from finance to economic growth.

Table 4. Results of the Hsiao causality tests

	Controlled variable	First manipulated variable	Second manipulated variable	FPE	*F*-statistics	ECM	Causality Inferences
Growth (G)	G (3)	lnF (1)		0.00051			
	G (3)	F (2)	C (1)	0.00050	6.489	-0.287**	C causes G
	G (3)	F (1)	M (1)	0.00024	11.562	-0.361*	M causes G
	G (3)	F (2)	PC(1)	0.00054	4.035	-0.0032	PC not cause G
Credit (C)	C (1)	F(1)		0.00552			
	C (1)	F(1)	G(2)	0.00364	6.332	-0.6156*	G causes C
Private credit (PC)	PC(1)	F(1)		0.00503			
	PC(1)	F(1)	G(3)	0.00278	5.729	-0.847*	G causes PC
M3(M)	M3(1)	F(1)		0.0029			
	M3(1)	F(1)	G(2)	0.0021	2.145	-0.8234*	G causes M3

Note: (*) and (**) denote significance at the 1% and 5% levels, respectively. The critical values are taken from the *t* distribution. Numbers in parentheses represent optimal lag lengths.

Figure 1. Causal Relationships



5. Discussions

The considerable evidence we find for bidirectional causality does not exclude the assumption dominant in the finance-growth nexus literature that finance leads economic growth. Both total credit and liquid liabilities predict economic growth in China while credit to private sector does not. However, the reverse causality from economic growth to private sector credit, which indicates that credit to private sector follows economic growth as a result of increased demand for financial services, does support Patrick's demand-following hypothesis. One possible interpretation of this evidence is that credit rationing is prevalent among Chinese private firms, which rely extensively on self-fundraising to meet their financing requirements.

Concisely, our results provide evidence for two-way causality between liquid liabilities and economic growth on the one hand and aggregate credit and economic growth on the other. Such bidirectional causality could mean that China's economic growth plays a key role by determining both the demand and supply sides of liquid liabilities and aggregate credit. Additionally, bidirectional causality between liquid liabilities and economic growth may suggest that the growth in total credit in China after 1978 played both a leading and accommodative role in economic growth. That is, by mobilizing savings generated by rising income, the banking sector in China succeeded in playing the critical role of recycler of financial resources, thereby further fueling economic growth.

Thus, endorsement of bidirectional causality would ensure the coherence and consistency of the Chinese finance-growth nexus and would add weight to the suggestion that the financial policies in China may not be as repressive as once thought. The results also support the idea that China's financial repression policies could rather be seen as financial restraint policies (see Hellmann et al, 1997) required for economic growth in the presence of asymmetric information and market failures. Thus, it appears that although the financial system has appeared inefficient at the microeconomic level, its contribution to growth via TFP may have been effective.

Despite the implications of significant levels of bad loan of Chinese financial institutions, it should be noted that a relatively large proportion of lending was for long-term fixed assets. In the traditional system, with only a few minor exceptions, banks did not lend for fixed investment. Long-term lending is inevitably more risky than short-term lending (Naughton, 1998). Bad loan was the unavoidable cost and a development option opted by Chinese authorities. From this reason,_in the past, Chinese banks were bailed out by the government after accumulating large amount of bad loans, which means that for state owned banks bad loan ultimately remains a public budget matter.

Considering the role of FDI (controlled variable in our VAR model) in the growth process of China, it can be argued that not only does the FDI trickledown effect depend on the extent the financial sector's development, but higher productivity is only possible when the host country has a minimum threshold stock of human capital (Alfaro et al., 2004). Human capital in turn is supported directly via formal education, on the job training and job creation (particularly in rural and remote areas) as part of the overall development goals, which, until 1995, banks paid considerable attention to in determining the allocation of bank credit (OECD's Economic Survey of China, 2005). Hence, it can be argued that the forms of government involvement in regulating and otherwise influencing the financial system in China has determined the degree to which the latter made a positive contribution to growth.

To imply that the Chinese financial system is inefficient one needs to prove that economic growth in china has been achieved without the theoretical contribution of financial sector, which would suggest that finance doesn't matter in the growth process or that the finance-led-growth proposition does not apply in the case of China. Such a proposition would be harder to explain in light of basic facts on china's financial evolution in the last 30 years and in regard to a number of empirical studies suggest that productivity improvement accounts for a significant proportion of China's spectacular growth (e.g., World Bank, 1997; Maddison, 1998; Wang & Yao, 2003; Jeanneney & Liang, 2000).

Alternatively, one has to prove that China's repressed financial system did not negatively affect its growth process. Therefore, shedding light on the puzzle requires the exploration of reasons other than those commonly evoked in the literature. For example, one possible avenue may well be spillovers created by 'financial inversion' in which a bad state bank loan can result in positive externalities in other sectors. Accordingly, in China, a state-owned bank's ability to fund projects which may yield low returns in the short-run may well coincide with high returns in the long run as external economies are generated. Thus the reasons for state intervention as well as the performance of the financial systems can be seen as macroeconomically efficient within the context of China's growth process. So far it can be argued that despite its inherent weaknesses. Therefore, the overall macroeconomic performance of the Chinese economy has been efficiently nurtured by its financial system's influence on allocation of resources and productivity growth.

Two limitations worth noting should be kept in mind when interpreting our results. First, the smallness of our sample size and the preponderance of theoretical reasoning on the efficiency issue, conclusions must be stated hesitantly and with ample qualifications. Second, sufficiently detailed measures of financial development are not available in China. For instance, appropriate datasets of informal financial transactions in China have not been included in the present study although the extent of informal financial markets in China suggests they have played a significant role in the growth process and therefore should be included in finance-growth nexus study of this kind.

6. Policy implications and conclusions

This paper aimed at providing a coherent perspective for understanding the multifaceted puzzle of China's financial development. To this end, the analysis empirically tests competing finance-growth nexus hypotheses using Hsiao's version of the Granger noncausality test for China over the period 1980–2002. Relevant literature has often asserted that financial development contribute directly to economic growth. Our empirical investigation also shows that this is likely to be the case in China. Our results also point to the existence of reverse causality running from economic growth to financial development. In brief, the major empirical results support a complex set of bidirectional causalities between Chinese financial development proxies and economic growth.

The evidence of bidirectional causality implies the co-evolutionary character of financial development and economic growth, which implies that China's financial intermediation is consistent with the country's economic growth requirement and developmental goals. Overall, the findings indicate that, at the macroeconomic level, China's financial development is rather efficient in respect to the country's developmental goals. That is, the paradox between China's impressive economic growth and its inefficient financial intermediation is only apparent when the nation is considered in terms of its level and pattern of economic development. Our finding that total credit and economic growth influence each other could also be interpreted as a denial of the financial repression hypothesis in favor of a *financial restraint* argument. Which also suggests the China's financial system does not distort the optimal allocation of funds. Such an interpretation does not underestimate the urgent need for strengthening the Chinese financial system through appropriate policies.

It is also worth noting that financial development in China has been one of the set of factors which made it possible for the high sustained growth to take place. More emphasis should be placed on growth-enhancing policies rather than on a big-bang type of financial reforms that tends to focus on short-run and sectoral returns. Even though this latter remains important, reforms should not proceed at a faster pace than structural changes in the real sector or changes taking place in institutional settings.

The study findings have several policy implications. First, the finding that credit to private sector does not contribute in predicting economic growth clearly indicates the need for further strengthening private sector and market-reform in China. Chinese banks should significantly improve awareness of risk management in order to successfully pass the test after their market-oriented reforms. In this regard, it is important to mention that China has taken serious steps to strengthen its banking sector by cleaning it up of bad loans ahead of deregulation mandated by the World Trade Organization at the end of 2006. It needs to be stressed that it is inevitable that significant efforts should be made in the next few years to restructure the financial system.

Second, with the risk that economic growth in China may be severely overheated reverse causality also recommends that in order to slow economic growth, tightening credit with tools other than interest rates (such as tight monetary expansion) will be effective in China. Moreover, under China's rapid economic growth financial system faces a risk resulting from overcapacity in a number of industries. Close interaction between finance and real sector recommend that in order survive boom-bust cycles, commercial banks should learn to reduce their exposure to potential bad loans resulting from excessive lending. Again, improving risk management and strengthening regulatory framework should be kept on top of reform agenda in China.

Our conclusions must be taken moderately and with ample qualifications due to the small sample size and other limitations inherent shortcomings in finance-growth nexus. Chinese financial development is no exception to the long list of unresolved issues in the financial development literature. The Chinese case does, however, present a wide range of theoretical and empirical challenges; some of which this paper attempts to explain in light of its empirical findings.

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