Finance–Growth Nexus in China from an Endogenous Switching Perspective

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Abstract

This study examines the relationship between financial development and economic growth across Chinese provinces with switching causality. Four states are considered: bi-directional causality (state 1); one-way causality from growth to finance (state 2); one-way causality from finance to growth (state 3); and non-causality (state 4). While state 3 dominates in developed regions, states 1 and 3 occur intermittently in other regions. This implies that the demand for financial services induced by local economic growth plays a stronger role in driving financial development in under-developed regions. Consistent with prior research, bank loans negatively affect economic growth in China.

Keywords: Financial development; Economic growth; Switching causality; China regions.

JEL classifications: O16

1. Introduction

The relationship between financial and economic developments, which serves as a convenient benchmark for justifying Schumpeterian growth theory, is among the most researched topics in growth literature (Schumpeter, 1911). The growth–finance nexus stems from two distinct views on such relationship (see, for example, Patrick, 1966; Levine, 1997).

First, increase in the demand for financial services as a result of economic growth may be the major driving force behind the development of the financial sector. Second, financial services may have a proactive role in promoting economic growth. Apart from the many original works that have proliferated, review articles on growth–finance also abound. For instance, Levine (2005) discussed the theoretical bases and empirical evidence of the finance–growth nexus. Most studies confirmed the presence of a linkage but the direction of causality lacks consensus. These inconsistencies are not surprising considering the spectrum of subject economies being studied and the variety of financial development proxies being used. This paper supplements the literature by exploring possible reversals in causality patterns.

Many studies have attempted to decipher causality patterns and found that finance-growth causality is likely to be associated with economic development (e.g., Calderón and Liu, 2003; Aghion *et al.*, 2005; Hassan *et al.*, 2011; Chow and Fung, 2013). Among them, the majority modeled causality as categorical events with no built-in reversal possibility. This practice is understandable because doing otherwise will require handling multiple dimensions instigated by the size of the variable/parameter set, the number of subjects, and the length of the time series. However, the directions of finance–growth causalities are unlikely to be time-invariant because the level of economic development influencing such directions are determined by time-varying socio-economic characteristics of the region, such as education, trade openness, and government policies. Performing Granger causality test in high-dimensional settings such as panel vector autoregressions (panel VAR) often takes the form of a simple equation-by-equation estimation, see e.g. Emirmahmutoglu and Kose, (2011). Dimension-reducing compromises, such as data grouping in the spirit of Frühwirth–Schnatter and Kaufmann (2008), must be incorporated even for exceptional cases like Chow and Fung (2013), whose framework permits switching in causality and system-wide estimation.

The present paper extends the framework of Chow and Fung (2013) by considering regime switching finance–growth causality with time-variant and potentially subject-dependent transition probabilities. The model is applied to regional data of China to examine the relationship between financial development and economic growth with possibilities of causality reversals. We focus on a single country, namely China, because cross-country growth regressions may not fully capture country-specific idiosyncrasies and are susceptible to data compatibility problems across countries (Levine and Zervos, 1996). Guariglia and Poncet (2008) and Hasan *et al*. (2009) argued that sub-national studies have major advantages over cross-country studies in addressing these issues. Studies of the growth–finance nexus using cross-regional data in a single country emerged as a result (e.g., Jayaratne and Strahan, 1996; Dehejia and Lleras–Muney, 2003; Koetter and Wedow, 2010; Guiso*et al.*, 2002).

Inasmuch as the finance-growth causality is associated with economic development, China is arguably an ideal avenue for investigating possible causality switches due to the sequential implementation of economic reform among Chinese provinces.[1] Following Guariglia and Poncet (2008), we collect provincial data from China because the extraordinary growth of this country is unlikely to be fully explained by standard growth regressions and the determinants of such growth are far from certain.

2. Literature Review

Two distinct views of the finance-growth nexus in development economics are: (1) the increase in the demand for financial services resulting from economic growth is the major

[1] Economic reform was first implemented in coastal provinces in 1979 and then sequentially extended to inland provinces in the late 1990s.

driving force behind the development of the financial sector (Robinson, 1952, p. 86), and; (2) financial services play a proactive role for in promoting economic growth (Schumpeter, 1911; Goldsmith, 1969; McKinnon, 1973; Shaw, 1973). Findings from empirical studies have been mixed so far. Studies like Jung (1986), Demetriades and Hussein (1996), and Luintel and Khan (1999) found considerable evidence for bi-directional causality and that uni-directional causality from finance to growth is generally weak. Nevertheless, several other studies like King and Levine (1993), Rousseau and Wachtel (1998), and Bell and Rousseau (2001) found that financial development leads economic growth.

Another strand of research documented that finance-growth causality is likely to be associated with economic development. Calderón and Liu (2003) validated the heterogeneity in the finance–growth linkage among subject economies and found that the contribution of finance hinges on the state of development and industrialization of the country concerned. Patrick (1966) argued that the pair is associated with the stage of economic development of the country in such a way that the mutually reinforcing relationship between the two would diminish with sustained economic growth. Supporting this view, Aghion *et al*. (2005) found a vanishing positive effect of financial development on steady-state per-capita GDP. This issue was also addressed by Hassan *et al.* (2011) and Chow and Fung (2013).

Due to country-specific idiosyncrasies and data compatibility issues that come with cross-country growth regressions, several studies used cross-regional data within a single country to examine the relationship between financial development and economic growth. For instance, Jayaratne and Strahan (1996) and Dehejia and Lleras–Muney (2003) used data from different states across the U.S.; Koetter and Wedow (2010) focused on the finance-growth nexus in regions of Germany; Guiso *et al.* (2002) performed the analysis using cross-regional data in Italy; and Zhang *et al* (2015) studied the impact of openness on China's financial development. Levine (1997, 2005), meanwhile, conducted comprehensive literature reviews on the finance-growth nexus.

Not surprisingly, findings concerning the links between growth and finance in China have been inconclusive as well. For instance, Liu and Li (2001) and Cheng and Degryse (2010) found a positive link between finance and growth in China. However, Aziz and Duenwald (2002) and Lu and Yao (2004) concluded that financial development has no impact on China's growth. Findings from Boyreau-Debray's (2003), Guariglia and Poncet (2008), Hasan *et al.*, (2009), and Lin *et al.*, (2015) even demonstrated a negative impact of financial deepening on China's growth due to the inefficient state-banking sector. In sum, regional variations in the growth-finance relationship seems to be a general phenomenon in China (e.g., Liang, 2005) and changes in causality patterns are worthy of further exploration.

3. Methodology

3.1 Core Panel VAR

Chow and Fung (2013) studied the finance–growth causality of 69 countries using a Markov switching panel VAR specification. The Granger causality between the two major variables at any time point could take on one of the following configurations: (i) one-way causality from finance to growth, (ii) one-way causality from growth to finance, (iii) bi-directional causality, and (iv) no causality. When imposed in the settings of Warne (2000) and Psaradakis *et al.* (2005), each of these configurations is associated with a state that can be realized by a Markov process. Back-to-back switching in causality or the continuation of a particular pattern are both possible depending on the data and the time-invariant transition probabilities[2].

[2] In Chow and Fung (2013), poorer countries and their more advanced counterparts have different causality patterns that closely map their differences in terms of financial openness and geographical proximity.

As shown in Canova and Ciccarelli (2009, 2013), estimating a full-fledged unrestricted panel VAR can be a formidable task even without the switching mechanism. Some restrictions are usually imposed on the cross-sectional or dynamic interdependence to alleviate the over-fitting problem and achieve estimation efficiency. Available options include pooling of cross-sections via clustering (Frühwirth–Schnatter and Kaufmann, 2008), dimension-reducing coefficient factorization (Canova and Ciccarelli, 2009), and specification of Bayesian priors with inherent cross-unit independence (Billio *et al.*, 2013). Chow and Fung (2013) used clustering approach and assumed the transition probabilities as exogenous, but the present study adopt the multivariate probit extension of Kim *et al.* (2008). The salient feature of this approach is that the transition probability of causality pattern within a subject region is no longer exogenous but is determined by a probit function that embodies correlations with the panel VAR errors and hence permits endogenous switching. Thus, switching is no longer an unexplained anomaly that is purely model induced, but is the result of changes in socio-economic factors and potential feedbacks of growth and financial development. Depending on the value of the parameters, our specification encompasses the usual time-homogeneous Markov switching process.

The basic unrestricted panel VAR model takes the following form:

$$y_{it} \equiv \begin{bmatrix} g_{it} \\ d_{it} \end{bmatrix} = \sum_{j=1}^{N} \sum_{l=1}^{p} b_{itl}^{j} \cdot y_{jt-l} + c_{it} \cdot v_t + u_{it}, \qquad (1)$$

where y_{it} is a $G \times 1$ endogenous vector that contains the economic growth g_{it} of region i at time t and the corresponding financial development indicator d_{it}.

Obviously, $G = 2$ in our context. The subject (unit) index runs from $i = 1, \cdots, N$, the time index runs from $t = 1, \cdots, T$, and p denotes the lag length of the panel VAR. Coefficient matrix b_{itl}^{j} has the dimension of $G \times G$; $v_t$is a $q \times 1$ vector of the exogenous and unit-independent variables, c_{it} is the $G \times q$ coefficient matrix, and u_{it} is the error term.

In this specification, the lag terms of the peer regions enter into the equation of unit i to reflect "dynamic interdependencies" as referred to by Canova and Ciccarelli (2013). The setup has an inherent time-varying feature because the parameters are allowed to differ across periods. The Granger causality (uni- or bi-directional) is statistically equivalent to restricting the off-diagonal elements of b_{itl}^{j}, for all j and l, to zero or non-zero[3].

Next, we briefly discuss our data sample before examining the imposed restrictions in (1). Our data set was taken from *China Data Online*, which was compiled by University of Michigan with data collected from China Statistical Year Book. As in prior research, we measure a region's level of financial development by bank loans. The data on province-level bank loans from China Statistical Year Book, which indicates the amount of bank loans extended by bank branches to their clients located in the same province, were collected by provincial offices of the China Banking Regulatory Commission. Banking institutions remained the dominant player in China's credit market during our study's sample period (1978-2011) and the emergence of non-bank financial institutions is a relatively recent event (see, for example, Cheng and Degryse, 2010). Fu and Heffernan (2007), for instance, documented that major banking institutions including policy banks, state-owned-commercial banks, joint stock commercial banks and foreign banks in China had a combined market share of about 84% of total loans in 2002.

Focusing on the post-reform era, we extract data from 1978 to 2011, convert regional output per capita into annual growth rates, log-transform bank loans per capita, and calculate the three–year averages of all variables to reduce the effect of short-term noises[4]. The original data set contains no inflation data and the price figures are retrieved separately from

[3] Chow and Fung (2013) assumed that the coefficient matrices were independent of time and across units. Instead, the subjects are pooled and grouped under different clusters. The time-varying property of the model is introduced via state-dependent coefficient matrix elements, the numerical values of which are similar for all subjects in the same cluster.
[4] Of the 31 provinces, autonomous regions, and municipalities, Tibet is the only region with no pre-1990 economic output data. Thus, Tibet was dropped from the sample.

the National Bureau of Statistics official website. Regional consumer price indices are generally available except for a few selective cases where pre-1980 data are lacking. In the latter case, we compute inflation using the retail price index instead.

** Insert Figure 1 here **

By analyzing these data from the perspective of Equation (1), we develop a system with $G = 2, N = 30,$ and $T = 11,$ where each period corresponds to a non-overlapping three-year interval (both ends inclusive). Figure 1 plots the regional data for selected time periods. Given the large dimension of the model and our use of three-year averages, a lag length of $p = 1$ is chosen. The situation is better explained by the following exposition. For subject $i = 1$ before the introduction of parameter restrictions, Equation (1) can be expanded as follows:

$$g_{it} = b_{it1}^{i(1,1)} g_{it-1} + b_{it1}^{i(1,2)} d_{it-1} + b_{it1}^{2(1,1)} g_{2t-1} + b_{it1}^{2(1,2)} d_{2t-1} + \cdots + b_{it1}^{30(1,1)} g_{30t-1} + b_{it1}^{30(1,2)} d_{30t-1} + c_{it}^{(1)} v_t^{(1)} + u_{it}^{(1)} \text{and} \quad (2)$$

$$d_{it} = b_{it1}^{i(2,1)} g_{it-1} + b_{it1}^{i(2,2)} d_{it-1} + b_{it1}^{2(2,1)} g_{2t-1} + b_{it1}^{2(2,2)} d_{2t-1} + \cdots + b_{it1}^{30(2,1)} g_{30t-1} + b_{it1}^{30(2,2)} d_{30t-1} + c_{it}^{(2)} v_t^{(2)} + u_{it}^{(2)}. \quad (3)$$

$b_{it1}^{i(a,b)}, a, b = \{1,2\}$ denote the row and column elements of the $G \times G$ coefficient matrix b_{it1}^{i}. All numbers in parentheses are similarly defined according to the locations of other coefficients or parameters in the system.

3.2 Causality and Switching

Block exclusion is the most common approach for testing Granger causality. For instance, in the context of (2), only the coefficients $b_{it1}^{i(1,1)}$ and $c_{it}^{(1)}$ can be non-zero if financial development does not Granger cause economic growth. However, when such non-causality exists in the panel data, we must decide whether to allow the financial development indicators (and growth) of peer regions to affect a particular subject. This issue is rarely addressed in other applications of panel Granger causality test because the dynamic interdependencies are often ignored. Although spatial relationships may exist between the regions, we are not sure whether such conjectural associations can be translated into causality patterns and how they can be translated. We follow the literature by excluding parameters $b_{it1}^{j(a,b)}, \forall j \neq i, a, b = \{1,2\}$ and by introducing an intercept term in $v_t^{(a)} = 1, \forall t, a = \{1,2\}$. Following Chow and Fung (2013), (2) and (3) can be augmented by using state indicators that embody different causality patterns. After removing the superfluous indexing subscripts and superscripts, the system can be rewritten as follows:

$$g_{it} = \left[b_i^{(1,1)} + \beta_i^{(1)} S_{it}^{(1)}\right] g_{it-1} + \left[b_i^{(1,2)} S_{it}^{(1)}\right] d_{it-1} + \left[c_i^{(1)} + \delta_i^{(1)} S_{it}^{(1)}\right] + u_{it}^{(1)} \text{and} \qquad (2')$$

$$d_{it} = \left[b_i^{(2,1)} S_{it}^{(2)}\right] g_{it-1} + \left[b_i^{(2,2)} + \beta_i^{(2)} S_{it}^{(2)}\right] d_{it-1} + \left[c_i^{(2)} + \delta_i^{(2)} S_{it}^{(2)}\right] + u_{it}^{(2)}. \qquad (3')$$

State variables $S_{it}^{(1)}$ and $S_{it}^{(2)}$ take a value of 1 if equation-wise causality exits and 0 otherwise. The 2×2 configurations can be summarized by a single indicator:

$$\mathbf{S}_{it} = \begin{cases} 1 & if\ S_{it}^{(1)} = 1, \quad S_{it}^{(2)} = 1, \\ 2 & if\ S_{it}^{(1)} = 0, \quad S_{it}^{(2)} = 1, \\ 3 & if\ S_{it}^{(1)} = 1, \quad S_{it}^{(2)} = 0, \\ 4 & if\ S_{it}^{(1)} = 0, \quad S_{it}^{(2)} = 0. \end{cases} \qquad (4)$$

The four scenarios in (4) correspond to bi-directional causality, uni-directional causality from growth to financial development, uni-directional causality from financial development to growth, and no causality, respectively. To obtain a more general specification, the covariance matrix of $u_{it} = \left[u_{it}^{(1)}, u_{it}^{(2)}\right]'$ will be assumed regime-dependent or $u_{it} \sim N\left(0, \Omega_{\mathbf{S}_{it}} | \mathbf{S}_{it}\right)$. The transition between the four states can be modeled in a few ways. Chow and Fung (2013) considered time-homogeneous Markov switching, but logistic transition (e.g., Frühwirth–Schnatter and Kaufmann, 2008) and probit transition (e.g., Amisano and Fagan, 2013) are viable alternatives. We employ a probit transition to allow time-varying transition probabilities and accommodate endogenous regime changes à la

Kim *et al*. (2008).

Kim *et al*. (2008) imposed an ordered probit switching[5] mechanism where the unobserved latent variable is correlated with the error term of the core model. This latter dependence provides the basis for endogenous switching. For notational convenience, we decompose the error terms in (2′) and (3′) as follows:

$$u_{it} = \begin{bmatrix} u_{it}^{(1)} \\ u_{it}^{(2)} \end{bmatrix} = \Lambda_{\mathbf{S}_{it}} \epsilon_{it}, \tag{5}$$

where $\epsilon_{it} = \left[\epsilon_{it}^{(1)} \epsilon_{it}^{(2)}\right]' \sim N(0, I)$ and $\Lambda_{\mathbf{S}_{it}} \Lambda'_{\mathbf{S}_{it}} = \Omega_{\mathbf{S}_{it}}$. The transition probabilities of (4) are specified as follows:

$$\mathbf{S}_{it} = \begin{cases} 1 & \text{if} \quad -\infty \le \eta_{it} < x'_{it}\gamma_{1,\mathbf{S}_{it-1}} \\ 2 & \text{if} \quad x'_{it}\gamma_{1,\mathbf{S}_{it-1}} \le \eta_{it} < x'_{it}\gamma_{2,\mathbf{S}_{it-1}} \\ 3 & \text{if} \quad x'_{it}\gamma_{2,\mathbf{S}_{it-1}} \le \eta_{it} < x'_{it}\gamma_{3,\mathbf{S}_{it-1}} \\ 4 & \text{if} \quad x'_{it}\gamma_{3,\mathbf{S}_{it-1}} \le \eta_{it} < +\infty \end{cases} \tag{6}$$

$$Pr(\mathbf{S}_{it} = s | \mathbf{S}_{it-1} = r, x_{it}) = p_{rs}(x_{it}) = \Phi(c_{r,s,it}) - \Phi(c_{r,s-1,it}), \tag{7}$$

for $r, s = 1, \cdots, 4$, where$c_{r,0,it} = -\infty$; $c_{r,4,it} = +\infty$; $c_{r,k,it} = x'_{it}\gamma_{k,r}$ for $0 < k < 4$, $\Phi(\cdot)$ is the standard normal cumulative distribution function, and η_{it} is the latent variable of the probit process that is distributed as $N(0, 1)$.

Since previous studies have mostly confirmed that the relationship between economic growth and financial development is associated with the stage of economic development, our controlling factors in x_{it} comprise a set of region-specific socio-economic variables often regarded as correlated with economic development in the empirical literature (e.g., Barro, 1991; Mankiw *et al*., 1992; Easterly *et al*., 1997). These variables include the logarithm of student enrolment–teacher ratio to proxy for education level, the logarithm of exports–GDP ratio to capture the degree of openness, and the logarithm of industrial output

[5] An alternative option is to model the switching by way of a multinomial probit (e.g., McCulloch and Rossi, 1994; Geweke *et al*., 1997) instead of an ordered probit.

share of state-owned-enterprises to control for reform progress. We also include the rate of inflation and the logarithm of real GDP per capita, which are conductive to the finance–growth relationship as argued by Rousseau and Wachtel (2002) and Rioja *et al.* (2007). To control for the policy privileges enjoyed by some regions, we use a dummy variable that indicates if the region is a Special Economic Zone (SEZ) and/or a Coastal Development Area in a certain period (the dummy is equal to 1 if affirmative and 0 if otherwise). The Markov nature of the state transition is enforced via the set of coefficients $\gamma_{k,\mathrm{S}_{it-1}}$ for the aforementioned controlling factors, which help define the thresholds in (6).

The feedback between the panel VAR model $(2') - (3')$ and the probit switching $(6) - (7)$ is instigated via the following non-zero correlation between the processes:

$$\begin{bmatrix} \epsilon_{it} \\ \eta_{it} \end{bmatrix} \sim N\left(0, \begin{bmatrix} I_2 & \rho \\ \rho' & 1 \end{bmatrix}\right), \qquad (8)$$

where the elements of the correlation vector $\rho = [\rho_1, \rho_2]'$ are subject to the usual $(-1, 1)$ bounds. Given this additional structure, the conditional density $f\left(y_{it} | \mathbf{S}_{it} = s, \mathbf{S}_{it-1} = r, \Theta_{\setminus\eta}\right)$ of the overall model, which is needed repeatedly in the simulation, is no longer normal. Here, $\Theta_{\setminus\eta}$ denotes the set of pre-determined variables and other parameters excluding $= \{\eta_{it}\}_{\forall it}$. By factorizing $[\epsilon_{it}, \eta_{it}]$ into $A[\epsilon_{it}, \omega_{it}]'$ with A being the Cholesky decomposition of the covariance matrix in (8), the following relationship is established:

$$\eta_{it} = \rho' \epsilon_{it} + \left(\sqrt{1 - \rho_1^2 - \rho_2^2}\right) \omega_{it}. \qquad (9)$$

By noting that $f\left(y_{it} | \mathbf{S}_{it} = s, \mathbf{S}_{it-1} = r, \Theta_{\setminus\eta}\right) = f\left(y_{it} |\, c_{r,s-1,it} \leq \eta_{\mathrm{it}} < c_{r,s,it}, \Theta_{\setminus\eta}\right)$ and from (9), we establish the following:

$$f(y_{it}|\mathbf{S}_{it}=s,\mathbf{S}_{it-1}=r,\Theta_{\backslash\eta})$$

$$=\frac{\phi\left(\Lambda_s^{-1}(y_{it}-C_s-B_s y_{it-1})\right)}{|\Lambda_s|p_{rs}(x_{it})}\left[\Phi\left(\frac{c_{r,s,it}-\rho'\left(\Lambda_s^{-1}(y_{it}-C_s-B_s y_{it-1})\right)}{\sqrt{1-\rho_1^2-\rho_2^2}}\right)\right.$$

$$\left.-\Phi\left(\frac{c_{r,s-1,it}-\rho'\left(\Lambda_s^{-1}(y_{it}-C_s-B_s y_{it-1})\right)}{\sqrt{1-\rho_1^2-\rho_2^2}}\right)\right], \qquad (10)$$

where $C_s+B_s y_{it-1}$ is the compact notation for the RHS of $(2')$and $(3')$, C_s, B_s denote the conforming parameter vectors/matrices under $\mathbf{S}_{it}=s$, and ϕ is the bivariate standard normal density function. Under the special case of exogenous switching (i.e., $\rho_1=\rho_2=0$), Equation (10) collapses into the conventional normal kernel.

4. Simulation

The complete model defined by $(2'), (3')$, and $(6)-(8)$ is estimated via Markov Chain Monte Carlo (MCMC) methods. The prior distributions for the panel VAR part are similar to those in Chow and Fung (2013) and involve the initial observations y_{i0}, the coefficient vectors b, β, c, and δ, and the covariance matrices $\Omega_{\mathbf{S}_{it}}$. For the ordered probit part, priors will be needed for initial states $\mathbf{S}_{i0}$, the state-specific coefficient vector $\gamma_{k,r}$ for $0<k<4$, $r=1,\cdots,4$, and the correlation vector ρ. Latent variables η_{it} and state variables $\mathbf{S}_{it}$ should also be sampled in the process.

For clarity of presentation, we rewrite the stacked panel VAR model as follows:

$$Y = Z\theta + U, \quad \Rightarrow \quad \begin{bmatrix} Y_1 \\ \vdots \\ Y_t \\ \vdots \\ Y_T \end{bmatrix} = \begin{bmatrix} Z_{0,\mathbf{S}_1} \\ \vdots \\ Z_{t-1,\mathbf{S}_t} \\ \vdots \\ Z_{T-1,\mathbf{S}_T} \end{bmatrix} \theta + \begin{bmatrix} U_1 \\ \vdots \\ U_t \\ \vdots \\ U_T \end{bmatrix}, \qquad (11)$$

$$\text{where} \quad Y_t = \begin{bmatrix} g_{1t} \\ d_{1t} \\ \vdots \\ g_{Nt} \\ d_{Nt} \end{bmatrix}, \quad U_t = \begin{bmatrix} \Lambda_{\mathbf{S}_{1t}} \begin{bmatrix} \epsilon_{1t}^{(1)} \\ \epsilon_{1t}^{(2)} \end{bmatrix} \\ \vdots \\ \Lambda_{\mathbf{S}_{Nt}} \begin{bmatrix} \epsilon_{Nt}^{(1)} \\ \epsilon_{Nt}^{(2)} \end{bmatrix} \end{bmatrix},$$

$$\theta = \left[b_i^{(1,1)}, \beta_i^{(1)}, b_i^{(1,2)}, c_i^{(1)}, \delta_i^{(1)}, b_i^{(2,1)}, b_i^{(2,2)}, \beta_i^{(2)}, c_i^{(2)}, \delta_i^{(2)} \right]'_{i=1,\cdots,N}$$

$$Z_{t-1,\mathbf{S}_t} = \begin{bmatrix} g_{1t-1} & S_{1t}^{(1)} g_{1t-1} & S_{1t}^{(1)} d_{1t-1} & 1 & S_{1t}^{(1)} & & & & 0_{1\times(10N-5)} \\ 0_{1\times 5} & S_{1t}^{(2)} g_{1t-1} & d_{1t-1} & S_{1t}^{(2)} d_{1t-1} & 1 & S_{1t}^{(2)} & & & 0_{1\times(10N-10)} \\ & & & \ddots & & & & & \\ 0_{1\times(10N-10)} & g_{Nt-1} & S_{Nt}^{(1)} g_{Nt-1} & S_{Nt}^{(1)} d_{Nt-1} & & 1 & S_{Nt}^{(1)} & & 0_{1\times 5} \\ 0_{1\times(10N-5)} & & S_{Nt}^{(2)} g_{Nt-1} & d_{Nt-1} & & S_{Nt}^{(2)} d_{Nt-1} & 1 & & S_{Nt}^{(2)} \end{bmatrix}.$$

In this equation, θ is a $10N \times 1$ vector, Y_t and U_t are $2N \times 1$, and the matrix $Z_{t-1,\mathbf{S}_t}$ has a dimension of $2N \times 10N$. Given the correlation between ϵ_{it} and η_{it} as defined in (8), the likelihood of the panel VAR conditioned on $\eta = \{\eta_{it}\}_{\forall it}$ can be denoted as follows:

$$\mathcal{L}(\theta | Y, Z, \eta) \propto exp\left\{-\frac{1}{2}(Y - Z\theta - \xi)'\Psi^{-1}(Y - Z\theta - \xi)\right\}, \qquad (12)$$

where $\xi = \tilde{\eta} \otimes \rho$ is the conditional mean of ϵ under (8), with $\tilde{\eta}$ being a$NT \times 1$ vector that is formed by stacking η_{it} first by i and then by t, and $\Psi = \sum_{s=1}^{4} I_{NT}^{s} \otimes [\Lambda_s(1 - \rho'\rho)\Lambda_s']$, where I_{NT}^{s} is an $NT \times NT$ diagonal matrix with the it-th entry set to 1 if $\mathbf{S}_{it} = s$ and 0 if otherwise. The conditioning implies that explicit sampling of latent vector $\tilde{\eta}$ simplifies certain steps of the algorithm (e.g., Kang, 2014) to deal directly with the Gaussian kernel (12) instead of the skewed counterpart (10).

Upon initialization of the starting values, the MCMC program is iterated through the following steps:

1. Sampling $\{y_{i0}\}$: We use the same normal prior as in Chow and Fung (2013) before sampling via a Gibbs step.

2. Sampling $\{\theta\}$: Assuming a normal prior for the regime-independent coefficient vector θ, a straightforward sampling is performed from the posterior by using a Gibbs step.

3. Sampling $\{\Omega_{\mathbf{S}_{it}}\}$: Assuming a wishart prior for the regime-dependent VAR error variances, sampling from the conjugate posterior is standard.

4. Sampling $\{\eta_{it}\}$: The inherent thresholds in (6) and the conditional density (10) indicate that the latent variable can be sampled sequentially from the truncated normals after $\{\mathbf{S}_{it}\}$, $\{\Omega_{\mathbf{S}_{it}}\}$, and $\{\epsilon_{it}\}$ are known. Specifically, we draw the following:

$$\begin{aligned}&\eta_{it} \mid \mathbf{S}_{it} = s, \mathbf{S}_{it-1} = r \quad \sim \\ &N(\rho' \epsilon_{it}, 1 - \rho_1^2 - \rho_2^2) \cdot \mathbb{I}\left(c_{r,s-1,it} \leq \eta_{it} < c_{r,s,it}\right),\end{aligned} \tag{13}$$

for $r, s = 1, \cdots, 4$, where $\mathbb{I}(\cdot)$ is the indicator function that yields a value of 1 if the bracketed argument is true, and 0 if otherwise.

5. Sampling $\{\gamma_{k,r},\ 0 < k < 4,\ r = 1, \cdots, 4;\ \rho\}$: Given that the block of parameters is denoted by ψ, the full conditional of ψ is not a tractable form (Kang, 2014). From (10), the necessary draws are performed via a Metropolis step with the following kernel:

$$\prod_i \prod_t (d\Phi_t)^{\mathbb{I}(\mathbf{S}_{it-1}=r)}, \tag{14}$$

$$d\Phi_t = \begin{cases} \Phi\left(\dfrac{x_{it}'\gamma_{1,r} - \rho'\epsilon_{it}}{\sqrt{1-\rho_1^2-\rho_2^2}}\right), & \text{if} \quad \mathbf{S}_{it} = 1, \\ \Phi\left(\dfrac{x_{it}'\gamma_{2,r} - \rho'\epsilon_{it}}{\sqrt{1-\rho_1^2-\rho_2^2}}\right) - \Phi\left(\dfrac{x_{it}'\gamma_{1,r} - \rho'\epsilon_{it}}{\sqrt{1-\rho_1^2-\rho_2^2}}\right), & \text{if} \quad \mathbf{S}_{it} = 2, \\ \Phi\left(\dfrac{x_{it}'\gamma_{3,r} - \rho'\epsilon_{it}}{\sqrt{1-\rho_1^2-\rho_2^2}}\right) - \Phi\left(\dfrac{x_{it}'\gamma_{2,r} - \rho'\epsilon_{it}}{\sqrt{1-\rho_1^2-\rho_2^2}}\right), & \text{if} \quad \mathbf{S}_{it} = 3, \\ 1 - \Phi\left(\dfrac{x_{it}'\gamma_{3,r} - \rho'\epsilon_{it}}{\sqrt{1-\rho_1^2-\rho_2^2}}\right), & \text{if} \quad \mathbf{S}_{it} = 4. \end{cases}$$

6. Sampling $\{\mathbf{S}_{it}\}$: This step is achieved by using the forward filtering backward smoothing algorithm that is commonly used in state space models (e.g., Kim and Nelson, 1999; Kang, 2014). By suppressing miscellaneous parameters, the objective of the forward step is to update recursively the filtered probabilities for each subject i as follows:

$$f\left(\mathbf{S}_{it}|y_{it=1,\cdots,t};\cdot\right)$$
$$=\frac{\sum_{r=1}^{4} f\left(y_{it}|y_{it=1,\cdots,t-1},\mathbf{S}_{it},\mathbf{S}_{it-1};\cdot\right)f(\mathbf{S}_{it}|\mathbf{S}_{it-1}=r;\cdot)f\left(\mathbf{S}_{it-1}|y_{it=1,\cdots,t-1};\cdot\right)}{f\left(y_{it}|y_{it=1,\cdots,t-1};\cdot\right)}. \quad (15)$$

The first term in the numerator on the RHS is Equation (10), the second term is $p_{rs}(x_{it})$, and the third term is the updated probability from the previous loop. The denominator is the composite distribution of the numerator that is aggregated over all possible states r and s. After filtering, the smoothing part generates samples from $f\left(\mathbf{S}_{iT}|y_{it=1,\cdots,T};\cdot\right)$ and then backwards from the following:

$$f\left(\mathbf{S}_{it}=r|\mathbf{S}_{it+1},y_{it=1,\cdots,t+1};\cdot\right)$$
$$=\frac{f\left(\mathbf{S}_{it+1}|\mathbf{S}_{it}=r,y_{it=1,\cdots,t+1};\cdot\right)f\left(\mathbf{S}_{it}=r|\,y_{it=1,\cdots,t};\cdot\right)}{\sum_{r} f\left(\mathbf{S}_{it+1}|\mathbf{S}_{it}=r,y_{it=1,\cdots,t+1};\cdot\right)f\left(\mathbf{S}_{it}=r|\,y_{it=1,\cdots,t};\cdot\right)}, \quad (16)$$

where

$$f\left(\mathbf{S}_{it+1}|\mathbf{S}_{it}=r,y_{it=1,\cdots,t+1};\cdot\right)=\Phi\left(\frac{x'_{it+1}\gamma_{\mathbf{S}_{it+1},r}-\rho'\left(\Lambda^{-1}_{\mathbf{S}_{it+1}}\epsilon_{it+1}\right)}{\sqrt{1-\rho_1^2-\rho_2^2}}\right)$$
$$-\Phi\left(\frac{x'_{it+1}\gamma_{\mathbf{S}_{it+1}-1,r}-\rho'\left(\Lambda^{-1}_{\mathbf{S}_{it+1}}\epsilon_{it+1}\right)}{\sqrt{1-\rho_1^2-\rho_2^2}}\right). \quad (17)$$

5. Results

5.1 Overview of the Assessment Method

Tables 1 to 4 summarize the pooled MCMC output obtained from running 10 parallel chains of 5,000 iterations. The last 2,000 of each chain are used to compute the posterior estimates. Given the large model dimension, we only present the results for a selected subset of parameters. For the continuous parameters in Tables 3 and 4, we provide the posterior means with the numerical standard errors calculated using batch means. For categorical state variables $\mathbf{S}_{it}$ in Table 1, we evaluate the posterior mode as this satisfies the maximum-a-posteriori principle. Moreover, the configurations of the state variables in (4) are arbitrarily ordered, so assessing the posterior *medians* will not be appropriate as they are not invariant to a relabeling of the regime indices.

With regard to the overall validity of the model, we use the conditional predictive ordinate (CPO) of Gelfand (1996), which is measured by $f\left(y_i|Y_{\backslash i}\right)$. CPO is useful in performing "leave-one-out" cross-validation and allows us to assess how the model fits a particular observation. A higher CPO associated with the observation (the one being left out in the conditioning argument) indicates a better fit and vice versa. In addition, Gelfand (1996) showed that the CPOs could be pooled in a particular way to provide an aggregate indication of model fit. The pseudo Bayes factor (PsBF), which is the product of the CPO ratios that are defined in our context, is expressed as follows[6]:

$$PsBF = \prod_{i=1}^{N}\prod_{t=1}^{T}\frac{f\left(y_{it}|Y_{\backslash it},\mathcal{M}_0\right)}{f\left(y_{it}|Y_{\backslash it},\mathcal{M}_1\right)}, \qquad (18)$$

where $\mathcal{M}_0$ and $\mathcal{M}_1$denote the model being assessed and the alternative model, respectively. We compare the adequacy of our switching model with those models that have causality belonging to one of the four classes depicted in (4) but are time-invariant. A positive log PsBF supports the null model and a negative figure indicates the opposite.

5.2 Implications on Switching and Causality

Table 1 shows that our sample is dominated by state 3 (one-way causality from financial development to growth) and state 1 (two-way causality between growth and finance) with a

[6] Computationally, calculating the log transformed version, especially in high-dimensional problems, is more convenient than calculating the PsBF directly.

ratio of occurrence of about 63:37. Back-to-back switching is not pervasive, although switching *per se* is not rare. The average duration of a regime in any province is about 2.6 periods or nearly 8 years. So, the first important message from Table 1 is that non-causality is a non-issue empirically. Growth to finance causality never occurs in isolation, and when it emerges it is always accompanied by the reverse causality from finance to growth.

** Insert Table 1 here **

Some notable features can be found in the pattern of the realized regimes/states in the cross-section. First, financial development and economic growth are mutually reinforcing during the early stages of the economic reform which began in 1978. Bi-directional causality is observed in each region between 1979 and 1981, whereas one-way causality from finance to growth is not observed until 1981. Second, the economically developed regions (in terms of their 2011 per capita RGDP ranking which is the basis of the ordering of subjects in Table 1) experienced less switching in general. Among the top 10 developed regions which incidentally are all located along the coast, only Beijing and Zhejiang experienced back-to-back switchings. By contrast, similar phenomenon of high frequency switching occurred in eight provinces in the middle-10 tier and six in the bottom tier. Third, the developed regions are more prone to the dominance of finance to growth causality (state 3). Shanghai, Guangdong and Liaoning enjoyed uninterrupted finance to growth causality for ten periods, or 30 years from 1981 onward. Tianjin, Zhejiang, Inner Mongolia and Shandong each experienced state-3 causality for nine periods. On the other hand, the average number of periods classified as state 3 is less than six in the middle-10 tier.

Can the correlation between the economic growth and financial development in China be inferred under the presence of causality? To answer this question, we refer to Tables 2 and 3, which stipulate the posterior estimates of $b_i^{(a,b)},\ a,b = \{1,2\}$. Table 2 contains the coefficients of lagged financial development in the growth equation, which indicate the potential impact of financial development on economic growth in each region if either state 1 or 3 is realized in the regime switching process. The results show that two-thirds of the subject regions have recorded a negative value, which is consistent with the findings of other studies in which bank loans are negatively related to economic growth in China. Boyreau–Debray (2003), Guariglia and Poncet (2008), and Hasan *et al.* (2009) attributed this finding to high-level state interventionism, namely, the burden of supporting state-owned enterprises and the continuous poor lending practices in the banking sector of China. This argument was further elaborated and evidenced by Lin *et al*. (2015).

** Insert Tables 2 and 3 here **

Presented in a similar format to Table 2, Table 3 reports the coefficients of lagged growth in the financial development equation, which indicate the potential impact of economic growth on financial development in each region if either state 1 or 2 is realized. Differing from those in Table 2, the coefficients reported suggests a potential positive effect of growth on financial development.[7] This finding is consistent with the conventional view that financial development is driven by demand for financial services from the real sector.

Finally, we assess the validity of the observed causality pattern by testing it against possible alternatives. We compute the PsBF of our switching model ($\mathcal{M}_0$) against that of alternative models ($\mathcal{M}_1$) where a single state dominates and persists in all periods. Four alternatives are assessed. Each alternative corresponds to the static scenario of $\mathbf{S}_{it} = 1, 2, 3,$ or 4 only. The results are shown in the lower part of Table 2. Our specification beats all four alternative cases. In particular, state 4 (no Granger causality) and state 2 (one-way causality from growth to finance) have the weakest statistical evidence. To further cross-check our findings, we separately run a province by province bivariate VAR model and do

[7] As Table 1 indicates, such causality may not be realized in some regions due to the low probability for an independent realization of state 2.

the standard causality test. Granger causality tests conducted in such conventional way give no evidence of causality in either direction for nearly all subjects. Only two of the 30 provinces exhibit causality from growth to finance. Regardless of their statistical validity, these findings from the over-simplistic approach defy the general perception of most people regarding the growth history of China in the past few decades.

5.3 The Role of Switching Factors

This section attempts to link the finance–growth causality pattern discovered from the previous section with the switching factors considered in our study. Table 4 shows the posterior estimates of the selected coefficients of the probit part of our model and their corresponding marginal effects. The coefficients are relatively large because we imposed a non-informative prior. Such prior is less restrictive by design as no justifiable *a priori* assumptions are readily available for doing the opposite. Using more restrictive priors will hinder the switching potential of the model as it becomes impossible to differentiate underlying causality patterns even if they exist.

The so-called marginal effect, rather than the coefficient *per se*, matters the most in probit models. This effect refers to the change in the probability of a certain regime/state induced by a change in the switching factor. In our context, the marginal effect is the partial derivative of (14) with respect to the change in a particular continuous variable in x_{it}. For the dummy variable that indicates the presence of SEZ, we evaluate the discrete change in (14) for the dichotomous cases of with and without the SEZ status. In Table 4, we indicate the signs of the marginal effects instead of their values for computational reasons[8]. Given that the dominating regimes are states 1 and 3, we present only those situations where the previous state is either 1 or 3. The sign for some cases is uncertain because the standard

[8] For instance, we know that $\Phi(100) > \Phi(99)$, but they are both equal to 1 computationally. Therefore, the value of the marginal effect may not be correctly obtained.

normal pdf or cdf being evaluated depends on the sign and size of the probit coefficients and the integrated values cannot be ascertained numerically.

** Insert Table 4 here **

The table shows that three major indicators of economic development, namely, education (LOGST), inflation (INFL), and real per-capita GDP (LOGYCAP), consistently exhibit positive marginal effects on the persistence of or switching to state 1 (i.e., bi-directional causality between growth and finance) and state 3 (i.e., one-way causality from finance to growth). This comes in stark contrast to the findings of Patrick (1966), who proposed a diminishing relationship between economic growth and financial development, and of Aghion *et al*. (2005), who found a vanishing effect of financial development on per-capita GDP. On the other hand, the same three economic development indicators consistently show negative marginal effects on the probability of switching to state 2 (i.e., one-way causality from growth to finance) and state 4 (i.e., non-causality). In other words, non-causality and one-way causality from growth to finance are less likely to take place the more developed a region becomes, *ceteris paribus*.

5.4 Discussion of Major Findings

Taking the findings from Tables 1 and 2 together, while both states 1 are 3 are likely to persist in general, state 3 tends to be the dominating force in developed regions while states 1 and 3 occur intermittently in other regions. The dominance of state 3 in developed regions is in line with Stolbov's (2017) finding that growth-to-finance causality is a rare phenomenon while finance-to-growth causality still prevails among OECD countries. It is also consistent with Demirguc-Kunt *et al*.'s (2013) argument that economic progress boosts the development of securities markets (which provide customized financial arrangements) more than it boosts the development of banks (which provide standardized contracts).

The higher incidence of the growth-to-finance causality (as an element of state 1 but not an isolated event) in under-developed regions than in developed ones implies that demand for financial services created by local economic growth plays a stronger role in driving financial development in the former than it does in the latter. As Table 3 shows, such stimulating effect of economic growth on financial development is unambiguously positive and is more likely manifested by way of bi-directional causality in the less developed regions.

There are two plausible explanations for the above findings. First, on the demand side, a region's financial sector is hard to develop without sufficient demand for financial services generated by its local real sector (Patrick, 1966). As a result, financial development in under-developed regions is relatively more sensitive to and dependent on local economic growth and this explains why one-way finance-to-growth causality (state 3) in these regions is less frequently observed than in their more developed counterparts.

Second, from the perspective of credit supply, Boyreau-Debray and Wei (2005) found in China's banking sector a tendency for capital to move from high-productivity to low-productivity regions due to state polices. This external source of capital to under-developed regions reduces the extent to which capital supply is constrained by local savings (Boyreau-Debray and Wei, 2005), and enhances the sensitivity of bank credit supply to local needs, which may have triggered the growth-to-finance causality (as an element of state 1) in developing regions. Moreover, the observed intermittent switches of the finance-growth causality regime between states 1 and 3 at the province level could be caused by province-specific factors such as policy expenditure shocks and how far the local banking sector responds to funding demand in developing regions. In the developing Sichuan province, for instance, the switch from state 3 to state 1 during 2000-2002 could be explained by the large increase in demand for state-bank loans created by the large increase in policy expenditures on key infrastructure projects of highways, railways, airports and power plants (Sichuan

PDRC).[9] This demand-driven growth of the banking sector is captured by the growth-to-finance causality in state 1. Similarly, state 1 was realized again in Sichuan during 2006-2011 when policy expenditures on infrastructures were further enhanced. One can also find a similar phenomenon in the more developed Shandong province, where the regime of finance-growth causality switched from state 3 to state 1 in 2003-2005 due to the large increase in infrastructure investments by state-owned enterprises during this period (Shandong PDRC). Overall speaking, causality switches between states 1 and 3 are more frequently observed in low-income provinces where policy expenditure shocks were more frequent and state-banking was more pervasive relative to their high-income counterparts (Lin *et al*., 2012).[10]

Even though economic growth generates demand for financial services which serves as a necessary condition for under-developed regions' financial sector to grow in a sustainable way, the dominance of state 3 in developed regions conceals the subtle fact that growth could be hampered over the course of economic development. The finance-to-growth causality observed in China is partly the consequence of profound state interventions, such as the policy-oriented lending practices of state-owned banks. According to Guariglia and Poncet (2006), the large banking sector of China directs most bank loans to inefficient state enterprises than to efficient private enterprises.[11] The People's Bank of China estimated that 60% to 70% of non-performing loans in the banking system were caused by interventions and mandatory credit support of state-owned enterprises offered by the central and local governments (Goodstadt, 2014). Emergence of joint-stock commercial banks might have

[9] Each province's Provincial Development and Reform Commission (PDRC) formulates and implements plans for key project investments. Policy-oriented loans are borrowed through each province's local financing platform companies (Lu and Sun, 2013).

[10] According to Lin *et al*'s (2012) statistics, the big-4 state-owned banks' market shares were smaller in high-income provinces than in low-income ones.

[11] The distortions in the state-dominated financial sector force private enterprises to look for foreign investors (i.e., FDI).

improved the efficiency of China's financial sector (e.g., Fu and Heffernan, 2007), but their market shares remained small during the sample period. In 2002, for instance, the market share of policy and state-owned banks was 72% while that of joint-stock commercial banks was only 11%.[12] In sum, while economic growth is necessary for the financial sector to develop in the early stage, state intervention induces distortions and leads to the growth of an inefficient financial sector that in turn restrains economic growth. As Table 2 suggests, the growth of the financial sector has a negative effect on economic growth in two-thirds of the sample provinces. This result is consistent with the findings of previous research (e.g., Boyreau–Debray, 2003; Guariglia and Poncet, 2008; Hasan *et al*., 2009; Lin *et al*., 2015) and is arguably the main reason why feedbacks between finance and growth are more difficult to come by as a region becomes more developed.

6. Conclusion

This study investigates the relationship between financial development and economic growth among Chinese provinces. Previous studies on the causality between financial development and economic growth generally suggest two possible manifestations. First, the demand for financial services induced by economic growth drives financial development. Second, financial development has a proactive role in promoting economic growth by channeling funds to investment. Without ruling out either of these two possibilities, our empirical model allows the direction of causality to switch over time.

This study devises a regime-switching model for finance–growth causality where the transition probabilities are time-variant and are potentially subject-dependent. Four possible states (regimes) are considered, namely, bi-directional causality (state 1), one-way causality from growth to finance (state 2), one-way causality from finance to growth (state 3), and no

[12] Source: Almanac of China's Finance and Banking, 2003.

causality (state 4). The VAR model uses panel data and is augmented by a probit switching mechanism that allows us to evaluate endogenously the specific role of various socioeconomic factors in the economic development of China.

Consistent with previous studies, we find that bank loans, which is a typical measure of financial development, are negatively related to economic growth. This finding may be attributed to the adverse effect of state interventionism on the Chinese economy. We also find that typical economic development indicators, such as education, inflation, and per-capita GDP, increase the persistence of states 1 and 3, but decrease that of states 2 and 4. In particular, while state 3 is more frequently and persistently observed in prosperous regions, states 1 and 3 are observed intermittently in economically backward regions. This finding implies that the causality from growth to finance (as an element of state 1) is more likely to emerge in regions at their early stage of economic development. A ramification is that financial development in less developed provinces can be constrained at a relatively low level for some time due to the low demand for financial services from their real sectors. The underpinning of finance to growth is therefore less likely to be observed there. The prosperous regions may experience a restraint of another kind. The stimulation of finance to growth may not be a one-way process as market distortions induced by state interventions could hinder the pace of growth over the course of development.

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Table 1. Posterior Modes of the State Variables for all Provinces and Periods.

	1979-1981	1982-1984	1985-1987	1988-1990	1991-1993	1994-1996	1997-1999	2000-2002	2003-2005	2006-2008	2009-2011
Shanghai	1	3	3	3	3	3	3	3	3	3	3
Beijing	1	3	3	1	3	1	3	3	3	3	1
Tianjin	1	1	3	3	3	3	3	3	3	3	3
Jiangsu	1	1	1	3	3	3	3	1	1	1	3
Zhejiang	1	3	3	1	3	3	3	3	3	3	3
Inner Mongolia	1	1	3	3	3	3	3	3	3	3	3
Guangdong	1	3	3	3	3	3	3	3	3	3	3
Liaoning	1	3	3	3	3	3	3	3	3	3	3
Shandong	1	3	3	3	3	3	3	3	1	3	3
Fujian	1	1	1	1	1	1	1	1	1	1	3
Jilin	1	1	3	3	1	3	3	1	3	3	3
Hebei	1	3	3	3	3	3	3	1	3	3	1
Hubei	1	3	1	1	1	1	1	3	1	1	3
Chongqing	1	3	3	3	3	3	3	3	3	3	3
Heilongjiang	1	1	3	1	3	3	1	1	1	1	1
Shaanxi	1	1	1	1	1	3	1	3	1	1	1
Ningxia	1	3	3	1	3	3	3	1	3	3	3
Shanxi	1	3	3	1	3	3	3	3	3	3	3
Hunan	1	1	1	3	1	1	3	1	3	3	1
Xinjiang	1	1	3	3	1	3	1	1	3	1	3
Henan	1	3	3	3	3	3	3	3	3	3	3
Qinghai	1	3	3	3	3	3	3	3	3	3	3
Hainan	1	1	1	3	3	3	1	3	1	3	3
Sichuan	1	1	1	1	1	3	3	1	3	1	1
Jiangxi	1	1	3	1	3	3	1	3	1	1	3
Anhui	1	3	3	1	3	1	1	1	1	1	3
Guangxi	1	3	3	3	3	3	3	3	3	3	3
Gansu	1	1	1	3	3	3	1	3	1	3	1
Yunnan	1	3	3	3	3	3	3	1	3	3	3
Guizhou	1	1	3	3	3	3	3	3	3	3	3

Remarks: The provinces are ranked by their 2011 per capita RGDP in descending order. State 1 (shaded) implies bi-directional causality between economic growth and financial development. State 3 indicates one-way causality from financial development to growth.

Table 2. Coefficients of Lagged Development in the Growth Equation and the Posterior Estimates of Other Selected Parameters.

Parameters	Posterior Means	Numerical s.e.	Parameters	Posterior Means	Numerical s.e.
$b_i^{(1,2)}, \quad i =$	$(1.0e-04\ \times)$	$(1.0e-05\ \times)$	$b_i^{(1,2)}, \quad i =$	$(1.0e-04\ \times)$	$(1.0e-05\ \times)$
Beijing	-0.0089	(0.2582)	Henan	0.0090	(0.3752)
Tianjin	0.0080	(0.3035)	Hubei	-0.0790	(0.2886)
Hebei	-0.0569	(0.3471)	Hunan	0.0075	(0.3138)
Shanxi	-0.0459	(0.2071)	Guangdong	-0.0421	(0.3138)
Inner Mongolia	-0.2632	(0.8006)	Guangxi	-0.0644	(0.3290)
Liaoning	-0.0351	(0.1203)	Hainan	0.0111	(0.3176)
Jilin	-0.0303	(0.2692)	Chongqing	0.0203	(0.2561)
Heilongjiang	0.0770	(0.2769)	Sichuan	-0.0154	(0.3136)
Shanghai	-0.0630	(0.4508)	Guizhou	0.0263	(0.2370)
Jiangsu	-0.0374	(0.2704)	Yunnan	-0.0439	(0.2409)
Zhejiang	-0.0335	(0.2912)	Shaanxi	-0.0155	(0.1996)
Anhui	0.0306	(0.3246)	Gansu	-0.0432	(0.3494)
Fujian	0.0860	(0.4067)	Qinghai	-0.0509	(0.2102)
Jiangxi	0.0276	(0.3250)	Ningxia	-0.0248	(0.1820)
Shandong	-0.0879	(0.2725)	Xinjiang	-0.0119	(0.1677)
$\Lambda_{\mathbf{S}_{it}=1}$			$\Lambda_{\mathbf{S}_{it}=3}$		
First diagonal	0.2931	(0.0041)	First diagonal	0.2936	(0.0042)
Second diagonal	1.1772	(0.0580)	Second diagonal	1.1837	(0.0590)
Off-diagonal	0.0671	(0.0053)	Off-diagonal	0.0730	(0.0062)
ρ_1	0.0024	(0.0037)	ρ_2	-0.0008	(0.0037)
$\underline{PsBF}$			$\underline{PsBF}$		
$\mathbf{S}_{it} = 1, \forall i, \forall t$	189.2		$\mathbf{S}_{it} = 3\ \forall i, \forall t$	184.0	
$\mathbf{S}_{it} = 2, \forall i, \forall t$	426.4		$\mathbf{S}_{it} = 4, \forall i, \forall t$	426.6	

Table 3. Coefficients of Lagged Growth in the Financial Development Equation.

Parameters	Posterior Means	Numerical s.e.	Parameters	Posterior Means	Numerical s.e.
$b_i^{(2,1)}$, $i =$			$b_i^{(2,1)}$, $i =$		
Beijing	0.0175	(0.0020)	Henan	0.0081	(0.0017)
Tianjin	0.0056	(0.0015)	Hubei	0.0104	(0.0017)
Hebei	0.0186	(0.0017)	Hunan	0.0090	(0.0016)
Shanxi	0.0152	(0.0018)	Guangdong	0.0130	(0.0018)
Inner Mongolia	0.0136	(0.0011)	Guangxi	0.0072	(0.0015)
Liaoning	0.0188	(0.0031)	Hainan	0.0229	(0.0018)
Jilin	0.0087	(0.0021)	Chongqing	0.0165	(0.0019)
Heilongjiang	0.0103	(0.0015)	Sichuan	0.0082	(0.0017)
Shanghai	0.0105	(0.0016)	Guizhou	0.0128	(0.0023)
Jiangsu	0.0178	(0.0025)	Yunnan	0.0166	(0.0020)
Zhejiang	0.0158	(0.0019)	Shaanxi	0.0222	(0.0020)
Anhui	0.0084	(0.0015)	Gansu	0.0169	(0.0018)
Fujian	0.0059	(0.0013)	Qinghai	0.0278	(0.0023)
Jiangxi	0.0134	(0.0019)	Ningxia	0.0151	(0.0020)
Shandong	0.0055	(0.0016)	Xinjiang	0.0236	(0.0022)

Table 4. Selected Probit Estimates and Marginal Effects.

Variables, $j =$	Constant	LOGST	INFL	LOGXY	LOGSOE	LOGYCAP	DZONE
$\gamma_{1,s_{it-1}=1}$							
Posterior Means	870.3	642.9	824..2	−2696.1	139.1	3946.2	567.2
Numerical s.e.	(325.1)	(326.1)	(414.3)	(214.1)	(209.4)	(197.6)	(220.6)
$\gamma_{2,s_{it-1}=1}$							
Posterior Means	−700.1	−1289.5	−617.8	516.7	796.4	−6112.4	2012.3
Numerical s.e.	(258.1)	(441.0)	(332.4)	(349.1)	(200.2)	(236.9)	(329.5)
$\gamma_{3,s_{it-1}=1}$							
Posterior Means	1118.6	2953.2	3026.1	−846.5	−146.6	3822.1	307.4
Numerical s.e.	(195.8)	(431.2)	(328.1)	(401.4)	(535.4)	(287.4)	(211.3)
$\gamma_{1,s_{it-1}=3}$							
Posterior Means	1957.9	59.1	2140.4	142.2	1028.2	4842.8	−2933.5
Numerical s.e.	(330.8)	(361.2)	(328.3)	(280.3)	(418.6)	(247.6)	(355.3)
$\gamma_{2,s_{it-1}=3}$							
Posterior Means	−1945.8	−904.9	−1735.4	662.7	−1062.1	−4857.0	−929.7
Numerical s.e.	(290.5)	(354.7)	(345.1)	(359.9)	(335.9)	(286.2)	(200.3)
$\gamma_{3,s_{it-1}=3}$							
Posterior Means	−464.1	595.0	−1566.9	−334.4	−5125.5	3519.6	−2365.1
Numerical s.e.	(351.0)	(352.7)	(430.6)	(185.5)	(317.4)	(291.0)	(238.0)
(Current regime, previous regime)	Sign of Marginal Effects Based on Average Observations						
(1, 1)	n.a.	+	+	−	+	+	+
(3, 1)	n.a.	+	+	−	−	+	−
(1, 3)	n.a.	+	+	+	+	+	−
(3, 3)	n.a.	+	+/−	−	+/−	+	−
(2, 1)	n.a.	−	−	+	+/−	−	+/−
(4, 1)	n.a.	−	−	+	+	−	−
(2, 3)	n.a.	−	−	+/−	−	−	+/−
(4, 3)	n.a.	−	+	+	+	−	+

Remarks: (1) The variables in the probit equation are as follows: (i) LOGST = log student to teacher ratio, (ii) INFL = CPI inflation, (iii) LOGXY = log exports to GDP share, (iv) LOGSOE = log state-owned enterprises' share of industrial output, (v) LOGYCAP = log RGDP per capita and (iv) DZONE = time dummy that indicates whether the region is a SEZ or a coastal development area. (2) The marginal effects are evaluated using the average values of the exogenous variables and the standard normal pdf and cdf.

Figure 1. Provincial Data of Economic Growth and Log Loan per Capita (Three Year Averages).