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# FOREWORD

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# CONTENTS

# ABSTRACT

A method is proposed to estimate efficiency of aggregate investment in a transitional economy, using provincial panel data from the People's Republic of China (PRC) as an experimental case. Inefficiency is defined on the basis of disequilibrium investment. It is further decomposed into allocative and production inefficiency. Allocative inefficiency is related to policy/institutional factors. The main findings are: the PRC investment demand hardly responds to capital pricing signals, whereas it is strongly receptive to expansionary fiscal policies and interprovincial network effect. Once institutional factors are separated out, there are clear signs of increasing allocative efficiency and receding growth in regional investment disparity. The estimates on production efficiency are broadly in line with regional development.

## I. INTRODUCTION

Capital investment plays a key role in economic growth, especially in an economy where there is relatively abundant labor supply. However, excess investment occurs when economic growth lags behind investment growth, due to lack of accompanying growth in capital productivity. Such inefficiency of investment at an aggregate level used to plague centrally planned economies (CPEs) (see, for example, Begg et al. 1990). An interesting question arises whether economic reforms of former CPEs have alleviated excess investment, or whether investment efficiency has improved as the market grows and prevails during economic transition. In the present study, we propose a model to explain how much institutional factors associated with CPEs affect aggregate investment efficiency in a transitional economy and estimate the model using a panel data set from People's Republic of China (PRC).

Several phenomena stand out concerning fixed-asset investment in the PRC over the last decade (see Figures 1 and 2). First, fixed-asset investment has been growing faster than gross domestic product (GDP), with an average rate of around 14 percent per annum in real terms as against nearly 10 percent in real GDP during 1990-2002; capital formation has also risen in terms of its GDP composition, from around 25 percent in 1990 to 36 percent in 2000. Such an outgrowth in capital formation is obviously unsustainable in the long run, implying the possibility of excess investment or decreasing capital productivity. Second, growth in capital formation has been volatile. As investment bears high adjustment costs, such a volatile movement must have incurred very high social costs,[1] let alone its adverse effects on inflation and output growth stability. Thirdly, investment growth fell sharply in the mid-1990s leaving total bank savings exceeding total bank loans for the first time since 1950. It makes us wonder if this symbolizes the end of "investment hunger"[2] or persistent investment shortage. If so, should this imply that aggregate investment in the PRC has finally become mainly responsive to market conditions and thus more efficient than before? Notice, however, that the surplus in bank savings seems to have helped encouraging nationwide government deficit financing at both the central and the provincial levels. It is unclear how much efficiency improvement a mixture of government-induced and market-induced investment activities can make in comparison with the situation under CPEs. What is clear is that the recent concern over banking sector reforms in the PRC, especially over problems of bad debts, relates closely to the problems of excess investment and of underutilized capital in production.

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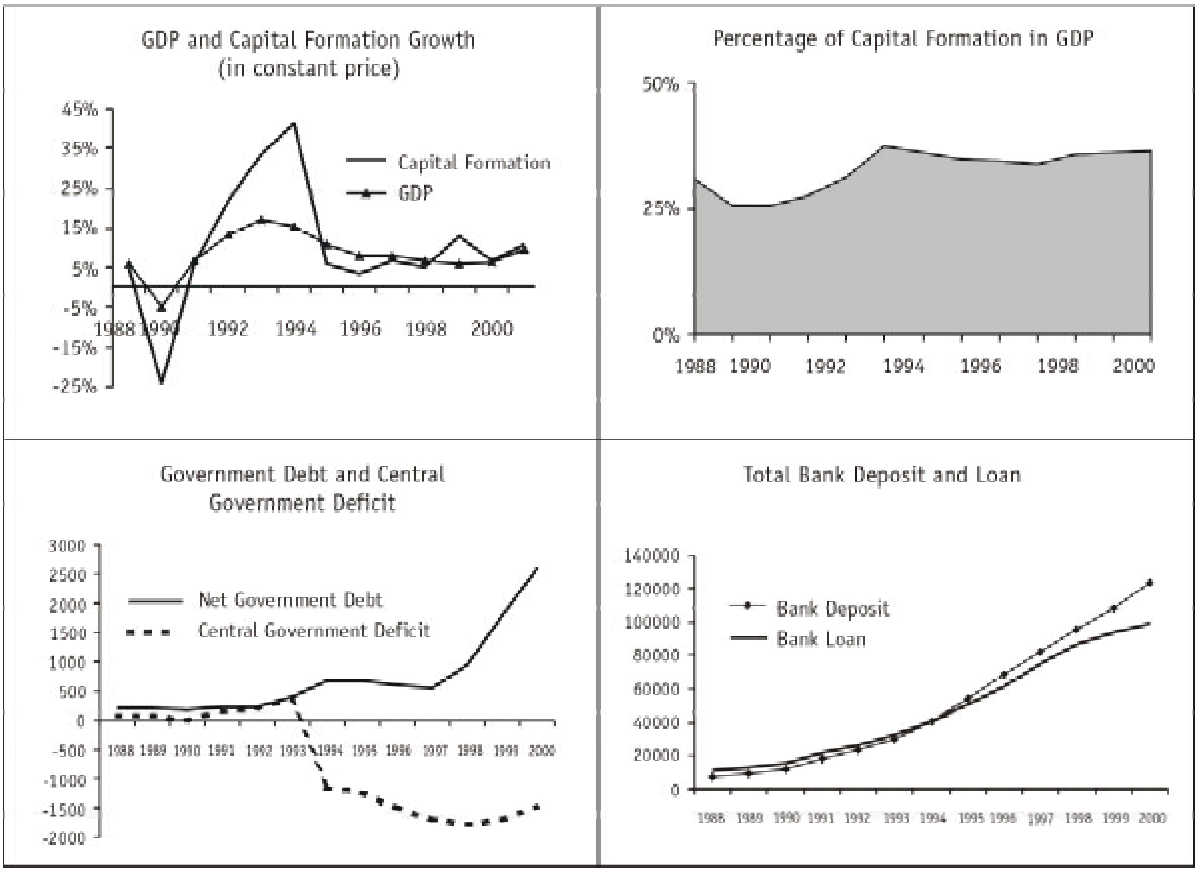
[1] This is implicitly confirmed by Sun (1998) and Song et al. (2001), who find that the PRC's short-run aggregate investment adjusts at very slow speeds to the long-run disequilibrium investment with respect to GDP.

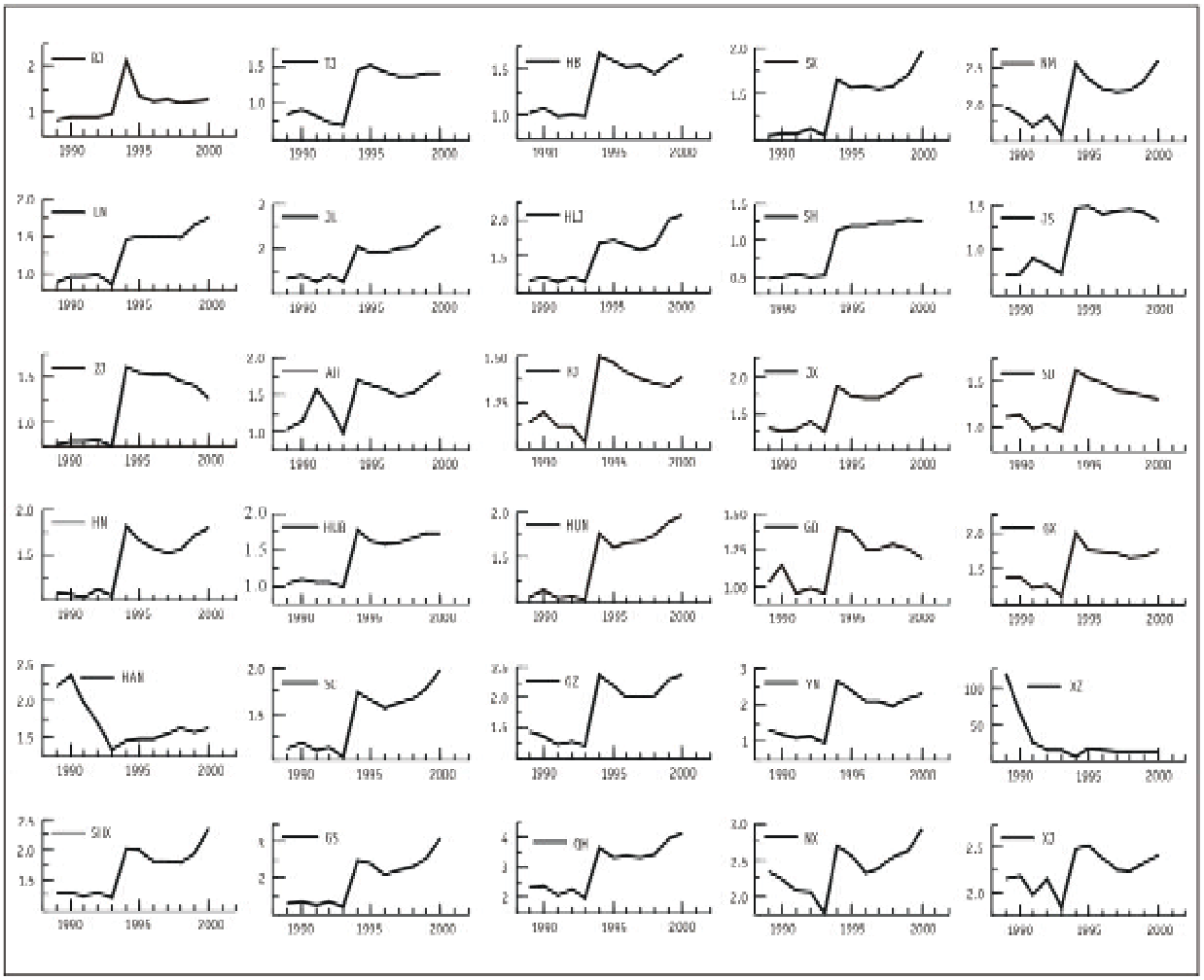
[2] Investment hunger is regarded as a key feature of CPEs, described initially by Kornai (1980).

**Figure 1: Capital Investment, GDP, and other Aggregate Series (in 100 million yuan)**




Recent studies on the PRC's aggregate investment lack conclusive views on the above questions. For example, Wang and Fan (2000) maintain that investment shortage is not yet over because of sizeable waste in past investment, which is reflected in unbalanced investment structure, policy-induced impulsive investment behavior, soft loans of the banking system, and relatively poor performance in a sizeable part of the state-owned sector. However, they reckon some signs of improvement in investment efficiency since the reforms, such as rising transformation rates from investment to capital formation, and increasing shares of investments by the nonstate-owned sector and the foreign sector. Zhang (2002) is very critical of the positive contribution of capital investment to the PRC's long-term growth. He regards investment outpacing GDP growth as a sign of excessive investment and of deterioration in investment efficiency. By showing decelarating growth in total factor productivity and diminishing investment returns during the 1990s, Zhang maintains that the PRC's overall investment in fixed assets has gone too far, especially with regard to its labor resource. He ascribes the problems mainly to institutional distortion, which induces a mixture of old tendency of excess investment with regional overcompetition for capital as a result of fiscal decentralization. The latter factor has attracted increasing attention in recent years. For instance, Zhang and Zou (1996) demonstrate empirically that a higher degree of fiscal decentralization is associated with lower provincial growth. They thus

**FIGURE 2: RATIO OF PROVINCIAL GOVERNMENT EXPENDITURE TO REVENUE**




Note: Due to the introduction of a new system of tax division in 1994, post-1994 data on local government revenue do not necessarily reflect the actual income of local governments, since the central government returns part of the tax collected nationally to provincial governments. Hence the above graphs can only represent trends of local government deficit financing rather than actual degrees of deficit.

infer that fiscal decentralization must have caused severe capital shortage for infrastructure investment at the national level, which is vital for rapid economic growth. The problem is more extensively examined by Young (2000), who demonstrates that decentralization has resulted in significant fragmentation of internal markets and therefore worsened efficient allocation of resources. But these empirical findings are somewhat at odds with Huang's detailed analysis of the political economy of central-local relations in investment controls (1996). Huang argues that the PRC's present de facto federal system, in which economic responsibilities are delegated to the local governments while the central government keeps political responsibilities, can have the merits of reducing coordination costs and improving economic governance. The economic role of federalism is further theorized by Qian and Roland (1998), who postulate two main effects. The first is competitive effect of federalism, which could lead to regional investment distortion; the second

is checks-and-balance effect of federalism, which should result in hardening soft budgets for state-owned firms. Unfortunately, these postulates lack rigorous empirical support.

In fact, there lacks systematic methods of measuring and evaluating efficiency of aggregate investment in the literature. This is reflected in a gap between the theoretical and empirical discussions on possible inefficiency in the PRC's aggregate investment. While theorists are most concerned with possible misallocation of financial resources due to imperfectly reformed economic systems,[3] empirical evidence is focused on production efficiency, such as productivity changes of capital in aggregation production functions or changing shares of capital to labor inputs. The problem, we believe, lies mainly in the different economic environments in which the issue has been considered. In a market economy, investment decisions are mostly made at the firm level and therefore the issue of investment efficiency falls formally in the realm of microeconomics; whereas in a transitional economy, the market is far from perfect and micro investment decisions are still significantly affected by various institutional factors.

The present study is an attempt to measure and evaluate inefficiency of aggregate investment in a transitional economy. We adapt capital input demand theory and measures of investment efficiency in standard microeconomics to the case of aggregate investment. We extend the theory and the measures to cover a transitional economy. We disentangle investment efficiency into two types: efficiency in investment allocation and efficiency in capital utilization during production. We are particularly interested in identifying and estimating how institutional factors have contributed to excess investment via inefficient investment allocation. We experiment with our model using data of 30 provinces in the PRC over the period 1989-2000.[4] The arrangement of the paper is as follows: Section II presents a general theoretical framework for defining and measuring investment inefficiency; Section III extends the framework to transitional economies where there exist important institutional factors affecting investment decision making; empirical model results are reported and analyzed in Section IV, and the final section concludes the paper.

## II. INVESTMENT INEFFICIENCY: A CONCEPTUAL FRAMEWORK

In this section a simple model base is set up for the purpose of estimating inefficiency in aggregate investment. The model is adapted from microeconomics. The problem of aggregation is disregarded for simplicity, following the normal practice of most of the empirical macro models of investment (see Caballero 1999). Our key focus is on how to measure inefficiency. We start by defining excess investment demand as deviations of actual investment from the desired investment driven by cost-minimizing factor demand for capital input. This enables us to exploit two available measures of efficiency—allocative efficiency and technical or production efficiency, and to relate investment deviations to these measures. The next section discusses the issue of how to link these measures to a transitional economy where there are serious institutional constraints to a perfect market situation.

[3] Bai et al. (1997) point out that improvement in production efficiency in terms of total factor productivity may not lead to more efficient resource allocation in a mixed market where firms are not solely profit maximizers.

[4] Beijing, Shanghai, and Tianjin are counted as provinces, but Chongqing, the new autonomous municipality, is still regarded as part of Sichuan.

Standard welfare economics dictates that perfect market equilibrium is the most efficient state. By this criterion, inefficiency in investment should arise largely from deviations of actual investment, $I$, from the market desired investment $I^*$. In the time-series context, we have:

$$\zeta_t = \ln I_t - \ln I^*_t \quad (1)$$

where $\zeta_t>0$ reflects excess investment and $\zeta_t<0$ under investment. Caballero et al. (1995) refer to $\zeta_t$ as "mandated" investment rate. We can regard it as disequilibrium investment rate if we define $I^*$ as equilibrium investment. The disequilibrium might be due to imperfect information, risk incurred because of uncertainty about the future, market imperfection, and decision-making errors. However, disequilibrium investment may not necessarily imply persistent inefficiency, though it is generally very costly to correct the existing disequilibrium toward equilibrium states. For the purpose of defining investment efficiency, we relate actual investment to capital $K$:

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

where $\delta$ is the effective depreciation rate for $K$. Defining $K^*$ as the long-run equilibrium capital stock, we should have:[5]

$$I^* = \delta K^* \quad (3)$$

Caballero et al. (1995) utilize the cointegration approach in order to measure of $\zeta_t$ in (1). Here, we choose to link investment directly to the production process via capital input demand. According to production theory, capital input is designated to be efficient if it is equal to the cost-minimizing factor demand for capital input under a given production process. The efficiency is further classified into two types. Production efficiency (PE)[6], which is associated with both the technological and managerial aspects of how capital assets are utilized in production, and allocative efficiency (AE), which concerns how production decisions are made in accordance with market demand and supply conditions (see, for example, Färe and Primont 1995; Greene 1997).

Let us consider a homothetic production function involving only two inputs—capital and labor:

$$Y = f(K, L, \Lambda) \quad (4)$$

where $\Lambda$ contains a measure of PE. Since we are mainly interested in long-run disequilibrium investment, we expect that the production function have constant returns to scale. Given (4), AE amounts to deviations from the equilibrium market condition of equality between the marginal rate of technical substitution between the inputs and their equilibrium price ratio:

$$\frac{\partial Y / \partial K}{\partial Y / \partial L} = \frac{P_k^*}{P_l^*} \quad (5)$$

[5] An easy way of deriving this is via the equilibrium correction model (ECM). From (2), we have: $\Delta I_t = \Delta K_t - (I - \delta K)_{t-1}$ where $\Delta$ denotes first difference and the term inside the brackets corresponds to the long-run equilibrium solution.

[6] We avoid the more commonly used term "technological efficiency" because of its lack of emphasis on the managerial side, which should be more important for the PRC firms during the reforms.

An AE measure can then be defined by ratio of the actual price ratio to the equilibrium price ratio:[7]

$$Z_{kl} = \frac{P_k / P_l}{P_k^* / P_l^*} = \frac{Z_k}{Z_l}, \quad Z_j = \frac{P_j}{P_j^*}, \qquad j = l, k \tag{6}$$

Obviously, full AE means $Z_{kl} = Z_k = Z_l = 1$. Notice that there are two aspects of price distortion in (6) namely own-price distortion and relative price distortion. Since all prices are relative, the own-price distortion can be seen as one factor price distortion with respect to the general price level. Notice also that $Z_{kl} = 1$ can be achieved when both labor and capital prices are distorted at the same rate, as it only reflects AE with respect to resource allocation between the two factors. In practice, $P_j^*$ is unobservable. *Z*s are thus often viewed as a set of parametric correction in input factor prices. The set can be estimated either directly from the secondary price space of firms' cost-minimizing function constrained by a production function, or indirectly from the primal goods space of firms' input demand function conditional on cost minimization by means of input distance function (see Atkinson and Cornwell 1994, Atkinson and Primont 2002).[8]

In the present context, we are only interested in $Z_I$ and/or $Z_k$. If we choose the primal goods space, the AE measure of $Z_I$ amounts to the disequilibrium $\zeta$ in (1):

$$Z_I = \frac{I}{I^*} \quad \Rightarrow \quad lnZ_I = lnI - lnI^* = \zeta \tag{7}$$

The equation reveals that the cointegration method can be used as an empirical AE measure.

Let us now assume a CES (constant elasticity of substitution) function for (4) with constant returns to scale under equilibrium:

$$Y^* = \Lambda[\alpha K^{*\rho} + (1-\alpha)L^{*\rho}]^{1/\rho} \qquad 0 \neq \rho = 1 - \frac{1}{\sigma} < 1 \tag{8}$$

where $\rho$ is the *substitution* parameter mapping into $\sigma$, the elasticity of substitution. The factor demand function for the long-run $K^*$ corresponding to (8) when it is subject to cost minimization, i.e., $\min(P_k k + P_l L)$, becomes:

$$K^* = \alpha^\sigma \Lambda^{-1} Y^* \left( \frac{P_y^*}{P_k^*} \right)^\sigma \tag{9}$$

where $P^*_y$ is the minimum unit cost of output (see, for example, Varian 1992, chapter 4). Combining (9) and (3) into (1), we get:

[7] The actual market price ratio is more frequently used in equation (5) in the empirical literature. Under that context, firm-specific shadow prices are employed in contrast with market prices, e.g., see Baños-Pino et al. (2001).

[8] A detailed explanation of duality of the two approaches can be found in Färe and Primont (1995).

$$\zeta_t = ln I_t - \left[A + \beta_1 ln Y_t + \beta_2 ln C_t\right]$$
$$\overset{\beta_1=1}{=} ln\left(\frac{I}{Y}\right)_t - \left[A + \beta_2 ln C_t\right], \qquad A = ln\delta + \sigma ln\alpha - ln\Lambda \tag{10}$$

where $\beta_1$ is the inverse of the returns to scale parameter and hence is expected to be unity, $\beta_2 = -\sigma < 0$, and $C$ denotes user cost of capital relative to output cost, the standard specification of which is:

$$C = \frac{P_k}{P_y} = \frac{(r+\delta)}{(1-\pi)}\frac{P_I}{P_y} \tag{11}$$

where $r$ is the real interest rate for investment loans and $\pi$ is the tax rate. In view of our panel data set of 30 provinces, we can rewrite (10) as:

$$\zeta_{it} = ln I_{it} - \left[A_i + \beta_1 ln Y_{it} + \beta_2 ln C_{it}\right] \qquad i = 1, \cdots, 30$$
$$\overset{\beta_1=1}{=} ln\left(\frac{I}{Y}\right)_{it} - \left[A_i + \beta_2 ln C_{it}\right], \qquad A_i = ln\delta_i + \sigma ln\alpha_i - ln\Lambda_i \tag{10'}$$

Equation (10′) presents us with a convenient vehicle to estimate both measures of efficiency. According to the established procedure (see Greene 1997), PE corresponds to the fixed individual effect $\Lambda_i$ in $A_i$ of (10). Equation (10) tells us that identification of $\Lambda_i$ depends on knowing $\alpha_i$, which have to be estimated via the production function (8) unless either $\sigma = 0$ or $\alpha_i = \alpha \quad \forall i$, provided we have data for $\delta_i$. We also need to consider the possibility of rapid technological progress in $\Lambda_i$. This is dealt with here via the following alternative specifications:

$$\Lambda_i = exp\{\gamma_0 t + \gamma_i\} \qquad \text{(common trend plus individual effect)} \tag{12a}$$

$$\Lambda_i = exp\{\gamma_{0t} + \gamma_i\} \qquad \text{(common random time effect plus individual effect)} \tag{12b}$$

As for AE, full efficiency indicates $\zeta_{it}=0$. Discernibly, $\zeta_{it}$ can be easily identified with the residual term derivable from regressing $\ln I_{it}$ on $\ln Y_{it}$ and $\ln C_{it}$. However, a key conceptual weakness of this identification is that any structural interpretation of the residual term entails substantially oversimplified assumptions, cf. Qin and Gilbert (2001). Since AE forms our major concern, we devote the next section to ways of improving this measure.

## III. ALLOCATIVE INEFFICIENCY AND INSTITUTIONAL CONSTRAINTS

As mentioned before, most of the concern over the PRC's excess investment demand relates to financial resource misallocation due to imperfect market environment. The theoretical framework of the previous section does not cover this concern explicitly. Here, we argue that there are two types of AE regarding investment demand in a transitional economy. One results from those institutional arrangements that distort pure market demand conditions for investment. The other is the usual

type due to firms' decision errors, assuming that their investment decisions are already conditioned upon an imperfect market environment. Obviously, $\zeta_{it}$ of (10) does not allow us to identify the two types, except probably for the case when the estimated $\beta_2 = 0$,[9] i.e., investment demand is insensitive to price signals. We can interpret this as the actual $C$ being significantly different from market-equilibrating $C^*$, and thus infer the presence of imperfect market.

In this section, we propose two ways of modifying $\zeta_{it}$. The first is to modify the cost function to incorporate in it market-disequilibrating institutional effects. It is commonly recognized that many state-owned firms have objectives other than profit maximization (see, for example, Liu 2001, Dong and Putterman 2002). For instance, ideological concern for spatial equality and defence consideration used to be among the key objectives in state investment plans, see Ma and Wei (1997). These objectives are hard to achieve unless budget constraints are soft. In other words, a mixed-goal objective-maximizing function should correspond to a cost-minimizing function mixed with soft budget constraints. A common route to incorporate these institutional features is disaggregation, i.e., to formulate a two-sector model with different behavioral rules for the state-owned sector and the nonstate-owned sector. However, this route may not fully reflect the fact that it is becoming harder to differentiate firms' economic behavior simply by ownership in the PRC, since many firms suffer from incompletely specified property rights, or have their ownership diversified due to the gradual privatization programme, not to mention the extra cost of data requirement. An alternative is to specify soft-budget constraints by the degree of their capacity to alleviate hard cost constraints at the aggregate level. We adopt this approach and attach a multiplicative term $\tau(x)$ to the standard cost function:

$$(P_k K + P_l L)\,\tau\,(x) \tag{13}$$

where $x$ denotes a set of disequilibrating soft budget indicators such that $\tau(0) = 1$. For practical purposes, we choose the exponential function:

$$\tau(\boldsymbol{x}) = exp\left\{\prod\nolimits_j x_j^{\tau_j}\right\} \tag{14}$$

Substituting (14) into (13) and minimizing it subject to (8), we arrive at the following alternative to equation (10'):

$$\zeta_{it}^{\tau} = ln\, I_{it} - \left[A_i + \beta_1\, ln Y_{it} + \beta_2\, ln C_{it} + \sum\nolimits_j \tau_j' x_{jit}\right] \tag{15}$$

The difference between (10') and (15) gives us an explicit AE measure caused by $\tau(x)$. This measure has the advantage of directly evaluating both the positive and negative contributions of the institutional factors toward AE. It brings empirical model results closer to testing theoretical postulates concerning efficiency and evolving institutions during reforms.[10]

The second modification is to adapt the original interpretation of AE to $\zeta_{it}^{\tau}$, i.e., to try to interpret it as a measure of allocative inefficiency due to firms' decision errors in investment demand, while their decisions are already conditioned on a mixed market situation, as described in (13).

[9] Notice that (9) collapses into a simple acceleration model when this happens, i.e., when $\sigma$=0.

[10] Theorization of efficiency and institutional changes is still in the making (see, e.g., Yao 2002), and desires better interactions with applied studies.

Two considerations guide our adaptation. One is that regression residuals are mixed with all sorts of misspecification and/or measurement errors. These should be filtered out before we attempt inferring it as decision errors. The other consideration relates to the dynamics of error-correcting adjustment. If a measure of AE turns out to follow a white-noise process, as is normally assumed of the residual term of a regression model, we would always come to the conclusion that there is virtually no persistent allocative inefficiency. An interesting AE measure should thus be expected to follow a stationary process, which encompasses, rather than is identical to, a white-noise process. In the context of investment demand, corrections of financial resource misallocation are expected to be rather slow because of very high adjustment costs. This implies that the AE measures are likely to exhibit significant autocorrelation, and leads us to exploit the separate specification of a static, error-correction component from an innovative, nonstructural error term in time-series econometrics. More specifically, we adopt the normal practice of specifying (15) into an autoregressive-distributed lag (ADL) model, denoting its residual term as $\upsilon_{it}$. We propose to filter $\upsilon_{it}$ out from $\zeta_{it}^{\tau}$ before interpreting it as an AE measure.

We are now in the position to define two measures of AE. One measure, $z^{\tau}$, captures the institutional aspect of AE and the other, $z^{m}$, the conventional AE due to nonoptimal firm decisions:

$$\begin{aligned} z_{it}^{\tau} &= \zeta_{it} - \zeta_{it}^{\tau} \\ z_{it}^{m} &= \zeta_{it}^{\tau} - \upsilon_{it} \end{aligned} \tag{16}$$

Before moving on to empirical modelling, we need to consider how to select $x$. Two general principles underlie the selection. These variables must embody institutional disequilibrating effects, and they must satisfy $\tau(0) = 1$. We take especially into consideration those factors that have been suggested repeatedly in the relevant literature, such as regional factors arising from decentralization. A number of indicators are constructed and experimented, which cover national and local government fiscal policies, interprovincial competition, changes in firms' debt-asset ratios and in bank loan-deposit ratios. Four variables have survived the selection experiment, one at the national level and the other three at the regional level. More precisely,

$x_1$ denotes the nationwide effect of deficit-financing fiscal policies, which is taken as logarithm of the net government debt including both the central and the local governments;

$x_{2i}$ represents the local government expansionary fiscal policy effect, which is taken as logarithm of the ratio of provincial government expenditure to revenue;[11]

$x_{3i}$ is designed to capture the tendency of over-investment due to provincial competition, in addition to what $x_{2i}$ captures, which is defined as one-period lagged deviation of provincial excess investment from its average regional level; and

$x_{4i}$ is aimed at capturing regional growth effect, which is defined as one-period lagged deviation of provincial per capita GDP from its average regional level.

[11] Notice that post-1994 data on $x_{2i}$ do not represent as drastic an increase in provincial government deficit as Figure 2 suggests. This is because a new system of tax division was introduced in 1994, which entails part of the tax collected nationally to be returned to provinces by a certain formula, whereas the published local government revenue account does not contain this part. Nevertheless, local government deficit financing is mainly responsible for the nationwide government debt, as shown in Figures 1 and 2.

[12] Here, we adopt the division of three broad regions by the National Bureau of Statistics; see also Song et al. (2001).

Detailed definition of these variables and the division of three regions[12] are given in the Appendix. The debt-asset ratios and bank loan-deposit ratios have turned out to be insignificant and hence dropped out. This can be explained by the facts that few firms use bank loans exclusively for fixed capital investment, that the available debt-asset ratio data only cover the period of 1993-1999, and that most of these banking related series fluctuate far less than those fiscal policy variables.

## IV. EMPIRICAL RESULTS

To estimate disequilibrium investment based on (10), we use the following model:

$$ln\, I_{it} = a_0 + a_i + \beta_1 ln Y_{it} + \beta_2 ln C_{it} + \zeta_{it} \tag{10a}$$

where $\alpha_i$ denotes individual effect and its various specifications are given in Table 1. The data sample covers 30 provinces of 12 years, 1989-2000. Essentially, (10a) is expected to constitute a homogeneous long-run equilibrating, possibly cointegrating, relation. $\zeta_{it}$ should therefore be a stationary and probably nonwhite-noise process. We have to choose appropriate estimation methods accordingly. Considering that all the time series involved in (10a) are likely to exhibit nonstationary properties, we use two estimation methods: feasible generalized least squares (FGLS) method directly on (10a) and the dynamic panel model estimation method of combined generalized method of moments (GMM) on a first-order ADL version of (10a): [13]

$$\begin{aligned} ln\, I_{it} &= a_0 + a_i + b_0 ln\, I_{it-1} + b_{10} ln Y_{it} + b_{11} ln Y_{it-1} + b_{20} ln C_{it} + b_{21} ln C_{it-1} + \upsilon_{it} \\ \beta_1 &= \frac{b_{10} + b_{11}}{1 - b_0} \\ \beta_2 &= \frac{b_{20} + b_{21}}{1 - b_0} \end{aligned} \tag{10a'}$$

In order to check if these long-run coefficients withstand the rapid changes in the economy, we also conduct two subsample estimations in addition to full-sample estimation. The main estimation results are reported in Table 1.[14] As expected, the residuals of static model (10a) show strong autocorrelation, suggesting a very slow disequilibrium correcting process, whereas the residuals of the dynamic model are serially uncorrelated.[15]

It is noticeable from Table 1 that there is strong evidence supporting the postulate of constant returns to scale, i.e., $\beta_1$=1. Unsurprisingly, the estimates of this parameter are sensitive to the

[13] The estimation is carried out by PcGive 10. We use ordinary least squares (OLS) residuals as the weights of the FGLS estimator. For the GMM method, we choose one-step estimator since residual heteroscedasticity should not be a significant problem once the individual effects have been filtered out; see Arellano and Bover (1995), and also Blundell and Bond (1998).

[14] Since some sample observations of the cost variable are negative because of large negative real interest rates, we shift the real interest rate net of the depreciation rate upward by adding one to the whole series before taking log transformation. This adjustment should only affect the magnitude of the constant term.

[15] The significant first-order serial correlation is an expected feature of the GMM method. See Doornik and Hendry (2001, Chapter 7) for details of the residual autocorrelation test.

specification of time effects, especially in the dynamic panel model, since the time series of both investment and output are heavily trended. The closeness of FGLS estimates to the GMM long-run estimates without time effects implies cointegration of both series at $\beta_1$=1, which corroborates the findings by Sun (1998) and Song et al. (2001). It is found that the time effects remain largely insignificant in the form of either a deterministic trend or random effect. We henceforth drop the time effect specification. Another noticeable result is the very low significance level in $\beta_2$ estimates. There are two possibilities. Either $\sigma$, the elasticity of substitution, is virtually zero, or the actual $C_{it}$ has not been perceived as cost-minimizing signals. We are inclined to the latter based on the observation that bank loan rates and investment prices remained rather low during excess investment peaks in the sample period. To further verify this possibility, we carry out two experiments. We first investigate whether there are different responses to different components in *C*. Then, we examine whether there are nonhomogeneous responses to these signals. For the first experiment, we take $\beta_1$=1 and separate the cost variable in (11) into three parts:

**Table 1: Main Estimation Results for (10a)**

| **FGLS** | $\alpha_i = \gamma_i + \gamma_0 t$ | | | $\alpha_i = \gamma_i$ | | | $\alpha_i = \gamma_i + \gamma_{0t}$ | | |
|---|---|---|---|---|---|---|---|---|---|
| | 89-2000 | 92-2000 | 94-2000 | 89-2000 | 92-2000 | 94-2000 | 89-2000 | 92-2000 | 94-2000 |
| $\beta_1$ | 1.17 (0.1002) | 0.847 (0.1706) | 0.7795 (0.2694) | 1.152 (0.0213) | 1.144 (0.0367) | 1.126 (0.0546) | 1.135 (0.1068) | 0.7546 (0.1922) | 0.831 (0.288) |
| $\beta_2$ | -0.186 (0.0976) | -0.202 (0.1131) | -0.178 (0.1091) | -0.191 (0.0942) | -0.155 (0.1106) | -0.192 (0.1088) | -0.092 (0.1909) | -0.034 (0.2056) | -0.223 (0.1985) |
| $\gamma_0$ | -0.002 (0.0111) | 0.0329 (0.0184) | 0.0335 (0.0255) | | | | | | |
| Joint test of $\gamma$i $\chi^2$(30) | 1858 [0.000] | 1486 [0.000] | 1643 [0.000] | 1928 [0.000] | 1479 [0.000] | 1767 [0.000] | 1911 [0.000] | 1486 [0.000] | 1754 [0.000] |
| Joint test of $\gamma_{0t}$ $\chi^2$(sample size) | | | | | | | 6.914 [0.806] | 7.853 [0.448] | 4.285 [0.638] |
| Test: no residual autocorrelation (AR) | | | | | | | | | |
| AR(1) N(0,1) | 14.80 [0.000] | 12.73 [0.000] | 7.073 [0.000] | 14.88 [0.000] | 12.91 [0.000] | 7.26 [0.000] | 15.30 [0.000] | 14.80 [0.000] | 2.353 [0.019] |
| AR(2) N(0,1) | 5.426 [0.000] | 3.643 [0.000] | 0.5536 [0.58] | 5.471 [0.000] | 3.755 [0.000] | 0.6647 [0.506] | 5.647 [0.000] | 5.426 [0.000] | -2.399 [0.016] |
| AR(3) N(0,1) | -1.311 [0.19] | -2.439 [0.015] | -3.803 [0.000] | -1.368 [0.171] | -2.28 [0.023] | -3.892 [0.000] | -1.329 [0.184] | -1.311 [0.19] | -4.506 [0.000] |
| AR(4) N(0,1) | -5.719 [0.000] | -5.745 [0.000] | -5.717 [0.001] | -5.804 [0.000] | -5.579 [0.000] | -5.873 [0.000] | -5.824 [0.000] | -5.719 [0.000] | -3.291 [0.001] |
| Estimated long-run coefficients of (10a') | | | | | | | | | |
| **GMM** | $\alpha_i = \gamma_i + \gamma_0 t$ | | | $\alpha_i = \gamma_i$ | | | $\alpha_i = \gamma_i + \gamma_{0t}$ | | |
| | 89-2000 | | | 89-2000 | 92-2000 | 94-2000 | 89-2000 | | |
| $\beta_1$ | -0.2445 (0.4627) | | | 1.2043 (0.0869) | 1.4556 (0.1707) | 0.9185 (0.3291) | 0.1673 (0.6621) | | |
| $\beta_2$ | -0.6557 (0.3865) | | | -0.177 (0.3797) | -0.724 (0.552) | 2.7155 (2.6075) | -0.9847 (0.5811) | | |
| $\gamma_0$ | 0.1647 (0.0514) | | | | | | | | |

Note: Standard errors in round brackets and *p* value in squared brackets.

$$ln\left(\frac{I}{Y}\right)_{it} = a_0 + a_i + \beta_{20}\, ln\left(r+\delta_i\right)_t + \beta_{21}\, ln\left(\frac{P_I}{P_Y}\right)_{it} + \beta_{22}\, ln\left(1-\pi_{it}\right) + \zeta_{it} \qquad (10b)$$

We expect that $\beta_{20}$ <0, $\beta_{21}$ <0, and $\beta_{22}$ >0. The main estimation results are reported in Table 2. Clearly, these coefficient estimates differ considerably. Real net interest rate shows no significance; the relative price variable shows little significance; and the tax rate variable is significant in full-sample estimation but becomes insignificant as we move to more recent subsamples. Since the interest rate variable is so insignificant, we only try to see whether the poor significance of the relative price variable is due to the restriction of homogeneous response by estimating the following model:

$$ln\left(\frac{I}{Y}\right)_{it} = a_0 + a_i + \sum_{i=1}^{30} \beta_{21i}\, ln\left(\frac{P_I}{P_Y}\right)_{it} + \beta_{22}\, ln(1-\pi)_{it} + \zeta_{it} \qquad (10c)$$

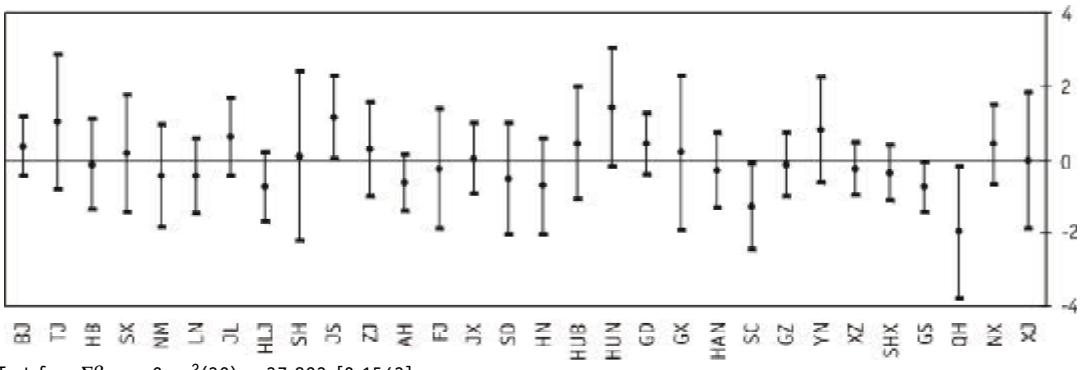
**Table 2: Estimated $\beta_2$'s in (10b)**

| | 89-2000 | | 92-2000 | | 94-2000 | | 89-2000 | |
|---|---|---|---|---|---|---|---|---|
| | FGLS | GMM | FGLS | GMM | FGLS | GMM | FGLS | GMM |
| $\beta_{20}$ | 0.0002<br>(0.0011) | 0.001<br>(0.003) | 0.001<br>(0.0012) | 0.0024<br>(0.0039) | -0.0012<br>(0.001) | 0.0003<br>(0.0036) | Restrict to 0 | |
| $\beta_{21}$ | -0.138<br>(0.1062) | -0.2484<br>(0.4641) | -0.1644<br>(0.1485) | 0.0182<br>(0.4027) | 0.5822<br>(0.2222) | 1.1285<br>(0.5871) | Restrict to 0 | |
| $\beta_{22}$ | 0.7953<br>(0.1658) | 0.9518<br>(0.3101) | 0.4199<br>(0.1933) | -0.6046<br>(0.441) | -0.072<br>(0.1919) | -0.0173<br>(0.354) | 0.7341<br>(0.1589) | 1.192<br>(0.2679) |

Note: Standard errors in round brackets and p value in squared brackets; GMM estimates are based on a first-order ADL of (10b).

Figure 3 plots the full-sample $\beta_{21i}$ estimates with 95 percent confidence intervals. Some heterogeneous response is apparent from the graph but the coefficients are insignificant overall, as shown from the test statistic reported below the graph. We henceforth drop out the interest rate and relative price variables.

**Figure 3: Price Coefficient Estimates for Individual Provinces (with 95 percent interval bars)**




Test for $\Sigma\beta_{21i} = 0$: $\chi^2(30) = 37.823$ [0.1543]

Next, we estimate the following version of (15):

$$ln\, I_{it} = a_0 + a_i + \beta_1\, lnY_{it} + \beta_{22}\, ln(1-\pi_{it}) + \tau_1' x_{1t} + \tau_2' x_{2it} + \tau_3' x_{3it\text{-}1} + \tau_4' x_{4it\text{-}1} + \zeta_{it}^{\tau} \qquad (15a)$$

as well as a first-order ADL of it similar to (10a′). Table 3 reports the main estimation results. The residual test results resemble closely those reported in Table 1.

It is evident from Table 3 that $\beta_1$=1 is strengthened. In other words, the slight tendency of decreasing return to scales, i.e., $\beta_1$>1 under $\alpha_i = \gamma_i$ in in Table 1 has disappeared and is very probably explained by one of the institutional variables. The estimates of $\beta_{22}$ fall and turn to wrong sign as we reduce sample size. This reinforces the earlier finding that investment demand hardly responds to capital price signals, implying that these signals are far from reflecting the real costs of investment. Notice that the combination of $\beta_1$=1 and $\beta_2$=0 enables us to simply view the ratio of investment to GDP as disequilibrium or "mandated" investment. Figure 4 plots the panel of this ratio. We see that Beijing, Shanghai, Gangdong, Hainan, and Tibet are among the most prominent for excess investment while Guizhou, Yunnan, and Guangxi are for underinvestment.

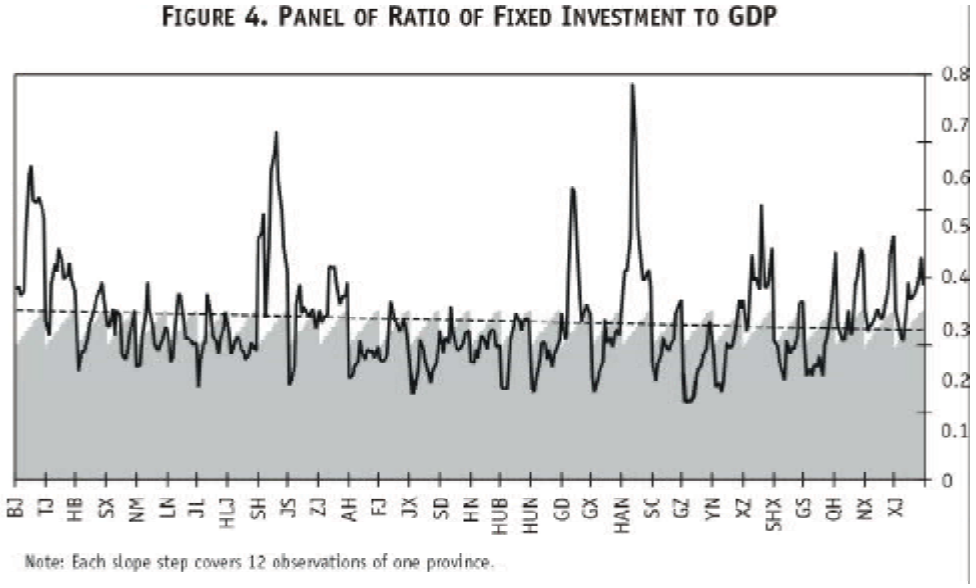
Now, let us look at the results of the institutional variables. Table 3 shows considerable differences between the FGLS and GMM coefficient estimates of these variables, and the latter estimates are mostly insignificant. This is because these variables are defined in rate, which differentiates them from flow variables in terms of time-series properties. In fact, the dynamic panel model estimation reveals that the way both $x_{i1}$ and $x_{i3}$ impact on investment are in first-order difference form. We thus respecify (15a) into a restricted dynamic model incorporating $\beta_1$=1 and $\beta_2$=0:

**Table 3: Main Estimation Results for (15a)**

| | 89-2000 | 92-2000 | 94-2000 | 89-2000 | 92-2000 | 94-2000 |
|---|---|---|---|---|---|---|
| | | FGLS | | | GMM | |
| $\beta_1$ | 0.9711<br>(0.0484) | 0.8782<br>(0.0603) | 0.9403<br>(0.0579) | 1.0378<br>(0.0719) | 1.0038<br>(0.1931) | 0.9114<br>(0.2316) |
| $\beta_{22}$ | 0.3603<br>(0.1427) | 0.2761<br>(0.1594) | -0.0805<br>(0.1614) | 0.0871<br>(0.4613) | -0.7739<br>(0.5455) | -1.0534<br>(0.616) |
| $\tau_1'$ | 0.0524<br>(0.0222) | 0.058<br>(0.0236) | 0.0459<br>(0.0208) | 0.0058<br>(0.0507) | -0.0245<br>(0.074) | -0.0208<br>(0.0772) |
| $\tau_2'$ | 0.0964<br>(0.0293) | 0.1973<br>(0.0433) | 0.1174<br>(0.0735) | 0.2129<br>(0.116) | 0.2985<br>(0.2268) | 0.3118<br>(0.2789) |
| $\tau_3'$ | -0.6727<br>(0.0698) | -0.6431<br>(0.0792) | -0.6308<br>(0.0755) | -0.0572<br>(0.2746) | 0.0755<br>(0.3748) | -0.1396<br>(0.2535) |
| $\tau_4'$ | 0.1556<br>(0.0729) | 0.1555<br>(0.0869) | 0.0195<br>(0.089) | 0.0941<br>(0.2809) | -0.2023<br>(0.5723) | -0.0269<br>(0.4244) |
| Test: No residual autocorrelation (AR) | | | | | | |
| AR(1)<br>N(0,1) | 9.693<br>[0.000] | 8.616<br>[0.000] | 6.479<br>[0.000] | -3.829<br>[0.000] | -3.481<br>[0.000] | -3.706<br>[0.000] |
| AR(2)<br>N(0,1) | 2.555<br>[0.011] | 2.235<br>[0.025] | 1.193<br>[0.233] | 0.9459<br>[0.344] | 1.679<br>[0.093] | 1.82<br>[0.069] |
| AR(3)<br>N(0,1) | -1.631<br>[0.103] | -1.702<br>[0.089] | -3.202<br>[0.001] | 0.9<br>[0.368] | 0.0661<br>[0.947] | -0.0616<br>[0.951] |
| AR(4)<br>N(0,1) | -4.821<br>[0.000] | -4.884<br>[0.000] | -5.804<br>[0.000] | -1.655<br>[0.098] | -0.5315<br>[0.595] | -0.959<br>[0.338] |

Note: Standard errors in round brackets and *p* value in squared brackets; GMM estimates are based on a first-order ADL of (15a).

**Figure 4. Panel of Ratio of Fixed Investment to GDP**




Note: Each slope step covers 12 observations of one province.

$$ln\left(\frac{I}{Y}\right)_{it} = a_0 + a_i + \tau'\Delta_1 x_{1t} + \tau_2' x_{2it} + \tau_3'\Delta x_{3it\text{-}1} + \tau_4' x_{4it\text{-}1} + \zeta_{it}^{\tau} \qquad (15b)$$

Table 4 reports the main estimation results. Model (15b) has the advantage of explicitly explaining disequilibrium investment exclusively in terms of institutional factors. Results from both models show that both fiscal policy variables have positively encouraged disequilibrium investment. Notice that $x_{i1}$ exerts its impact in a growth rate form. This suggests that changes in fiscal policies at the national level directly affect disequilibrium investment. As the rising government debt is due to deficit financing of many local governments, we infer that the positive impact of $x_{i2}$ helps to explain away the slight tendency of decreasing return to scales found in the estimates of model (10a). In other words, the part of persistent excess investment with respect to GDP, which leads to the inference of decreasing return to scale, can actually be accounted for by rising local government deficit spending. This suggests that investment induced by government deficit-financing policy is likely to encourage underutilization of capital, as judged by the expected long-run equilibrium $\beta_1$=1. The highly robust negative coefficient estimates for $x_{i3}$ are confirmatory of the view that provinces have been competing with each other to invest more if they notice that they have fallen behind their neighbors in the investment race. As for $x_{i4}$, its declining significance when we move to more recent subsample periods indicates that unequal regional allocation of investment due to unequal regional economic development has been gradually subsiding. Notice that this variable is somewhat negatively correlated with $x_{i2}$. This implies that provincial government expansionary investment policies may have contributed to the lessening of investment disparity to some extent, and that it is very difficult, if at all possible, for government to achieve efficiency and equality at the same time.

Let us now turn to the question of whether the institutional factors encourage allocative efficiency. First, we calculate $z^{\tau}$ of (16) using the following two residual series from full-sample FGLS estimation:

**Table 4: Main Estimation Results for (15b)**

| FGLS | 89-2000 | 92-2000 | 94-2000 | Correlation coefficients | | | |
|---|---|---|---|---|---|---|---|
| $\beta_1$ | Restrict to 1 | Restrict to 1 | Restrict to 1 | $\tau'_1$ | $\tau'_2$ | $\tau'_3$ | $\tau'_4$ |
| $\tau'_1$ | 0.0716<br>(0.0203) | 0.0466<br>(0.0202) | 0.0422<br>(0.0184) | 1 | 0.057 | 0.000 | 0.000 |
| $\tau'_2$ | 0.1463<br>(0.0266) | 0.2224<br>(0.0351) | 0.158<br>(0.075) | | 1 | -.033 | -.139 |
| $\tau'_3$ | -0.4831<br>(0.0827) | -0.5256<br>(0.0848) | -0.4678<br>(0.0828) | | | 1 | 0.005 |
| $\tau'_4$ | 0.2006<br>(0.0785) | 0.2065<br>(0.0901) | 0.145<br>(0.0956) | | | | 1 |
| Test: No residual autocorrelation (AR) | | | | | | | |
| AR(1)<br>N(0,1) | 13.11<br>[0.000] | 10.22<br>[0.000] | 7.103<br>[0.000] | | | | |
| AR(2)<br>N(0,1) | 6.414<br>[0.000] | 4.657<br>[0.000] | 1.35<br>[0.177] | | | | |
| AR(3)<br>N(0,1) | -0.6271<br>[0.531] | -0.4858<br>[0.627] | -3.433<br>[0.001] | | | | |
| AR(4)<br>N(0,1) | -5.606<br>[0.000] | -5.301<br>[0.000] | -6.293<br>[0.000] | | | | |

Note: Standard errors in round brackets and *p* value in squared brackets; GMM estimates are based on a first-order ADL of (15b) without restricting $\beta_1$=1.

$$\hat{\zeta}_{\text{it}} = ln\frac{I_{\text{it}}}{Y_{\text{it}}} - a_0 - a_{\text{i}}$$

$$\hat{\zeta}^{\tau}_{\text{it}} = ln\frac{I_{\text{it}}}{Y_{\text{it}}} - a_0 - a_{\text{i}} - \tau'_1\Delta x_{1\text{t}} - \tau'_2 x_{2\text{it}} - \tau'_3\Delta x_{3\text{it-1}} - \tau'_4 x_{4\text{it-1}} \qquad (16a)$$

Figure 5 plots the calculated $z^{\tau}$ for individual provinces, the cross-section means with 95 percent confidence interval bars over the sample period, and cross provincial covariance. Interestingly, most provinces show a rising $z^{\tau}$, and the rises are most prominent around 1993-1994 and in late 1990s when the PRC experienced major expansionary fiscal policy boosts. A slight rise is also discernible from the time series of cross-section means, notwithstanding the fact that both panel series of the two residuals in (16a) have zero means. The results show that institutional factors are likely to disencourage efficient allocation of investment, especially when a balanced fiscal policy is severely suppressed. Moreover, the dominance of positive over negative correlation between provincial $z^{\tau}$ in the covariance graph shows that provinces tend to suffer together from institution-induced allocative inefficiency, and that macro policy factors still exert great impact in the regional distribution of investment funds.

Next, we calculate $z^{\text{m}}$ of (16) using $\hat{\zeta}^{\tau}_{it}$ of (16a) and the residual series of the ADL model GMM full-sample estimation as $\hat{\upsilon}_{\text{it}}$ (column 5 in Table 3). The white-noise assumption of $\hat{\upsilon}_{\text{it}}$ is further confirmed by most of the normality test results at the provincial level reported in Table 5. Figure 6 plots the calculated $z^{\text{m}}$ for individual provinces, the cross-section means with 95 percent confidence interval bars over the sample period, and covariance between individual provinces.

FIGURE 5: $Z^{\tau}$ OF (16a), THEIR MEANS WITH 95 PERCENT CONFIDENCE INTERVAL BARS, AND COVARIANCE GRAPH

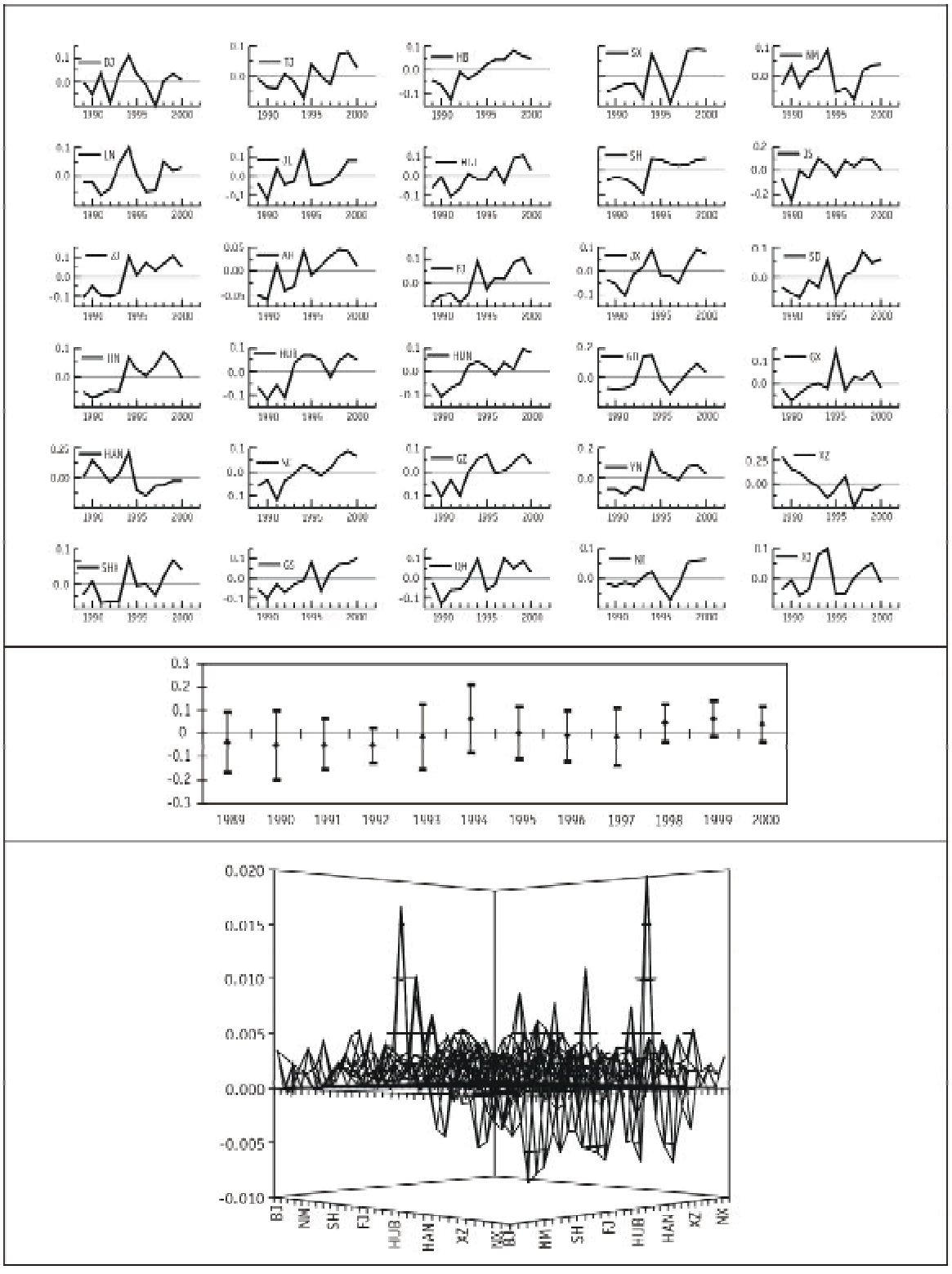


**Table 5: Normality Tests on $\hat{v}_{it}$ of (15a): $x^2$ (2)**

| BJ | TJ | HB | SX | NM | LN | JL | HLJ | SH | JS |
|---|---|---|---|---|---|---|---|---|---|
| 2.5475 [0.2798] | 5.911 [0.0521] | 3.5021 [0.1736] | 2.6254 [0.2691] | 0.8144 [0.6655] | 7.4409 [0.0242] | 3.6455 [0.1616] | 2.4106 [0.2996] | 6.6222 [0.0365] | 7.1045 [0.0287] |
| ZJ | AH | FJ | JX | SD | HN | HUB | HUN | GD | GX |
| 1.7466 [0.4176] | 4.0828 [0.1298] | 2.8665 [0.2385] | 0.5696 [0.7522] | 6.4875 [0.039] | 0.3909 [0.8225] | 8.6756 [0.0131] | 0.5158 [0.7727] | 2.766 [0.2508] | 2.5587 [0.2782] |
| HAN | SC | GZ | YN | XZ | SHX | GS | QH | NX | XJ |
| 7.0721 [0.0291] | 0.514 [0.7734] | 1.2515 [0.5349] | 0.6505 [0.7224] | 0.3539 [0.8378] | 6.3848 [0.0411] | 0.0637 [0.9686] | 0.0544 [0.9732] | 2.1116 [0.3479] | 1.7747 [0.4117] |

We see from the graphs that, in contrast to $z^{\tau}$, there is a trend of $z^{m}$ moving toward zero for many provinces, suggesting a general improvement of allocative efficiency in firms' aggregate investment demand as reforms proceed and the institutional effects have been filtered out. There is also a noteworthy contrast between the part of $z^{m}$ of around 1993-1994 and the part in the late 1990s for many provinces. While we can see a strong policy impact in the first part, i.e., firms' AE went up with that of the institution-induced AE around 1993-1994, the policy impact of the second part is hardly discernible from firms' AE in the late 1990s. The provinces with deteriorating firms' AE in late 1990s are concentrated in the less developed western and central regions (see Figure 4). All these demonstrate significant progress of decentralization and enhancing market conditions. The time series of cross-section means remain around zero and the covariance between provinces are more evenly distributed around zero, indicating that some provinces improve their firms' AE together while others are squeezed out by competition.

Notice that the contrasting trends in the two AE indices can be viewed as strong confirmatory evidence for the competitive effect and the checks-and-balance effect postulated by Qian and Roland (1998). To facilitate further comparison, we have calculated various rank correlation coefficients of the two AE measures (see Table 6). It is interesting to see from the rank autocorrelation coefficients that evolution of the institutional AE measure follows closely macro changes in the PRC's political economy whereas improvement of firm-level AE is more gradual and persistent. Correlation between $z^{\tau}$ and $z^{m}$ over time shows certain sign of disassociation, suggesting that firms' investment decisions have become less affected by institutional considerations as reforms deepen.

Finally, we calculate three versions of PE using the following equation based on (10):[16]

$$\begin{aligned}\Lambda_i &= \frac{\bar{\delta}_i \alpha_i^{\sigma}\, exp\{-a_i\}}{max\{\Lambda_i\}} \\ &\overset{\bar{\delta}_i=\bar{\delta}}{=} \frac{\alpha_i^{\sigma}\, exp\{-a_i\}}{max\{\Lambda_i\}} \\ &\overset{\sigma=0}{=} \frac{\bar{\delta}_i\, exp\{-a_i\}}{max\{\Lambda_i\}}\end{aligned} \qquad (17)$$

[16] Most of the PE indices use the negative of the fixed individual effects to reflect the degree of technological inefficiency. Our indices denote PE directly.

FIGURE 6: $z^m$, THEIR MEANS WITH 95 PERCENT CONFIDENCE INTERVAL BARS, AND COVARIANCE GRAPH

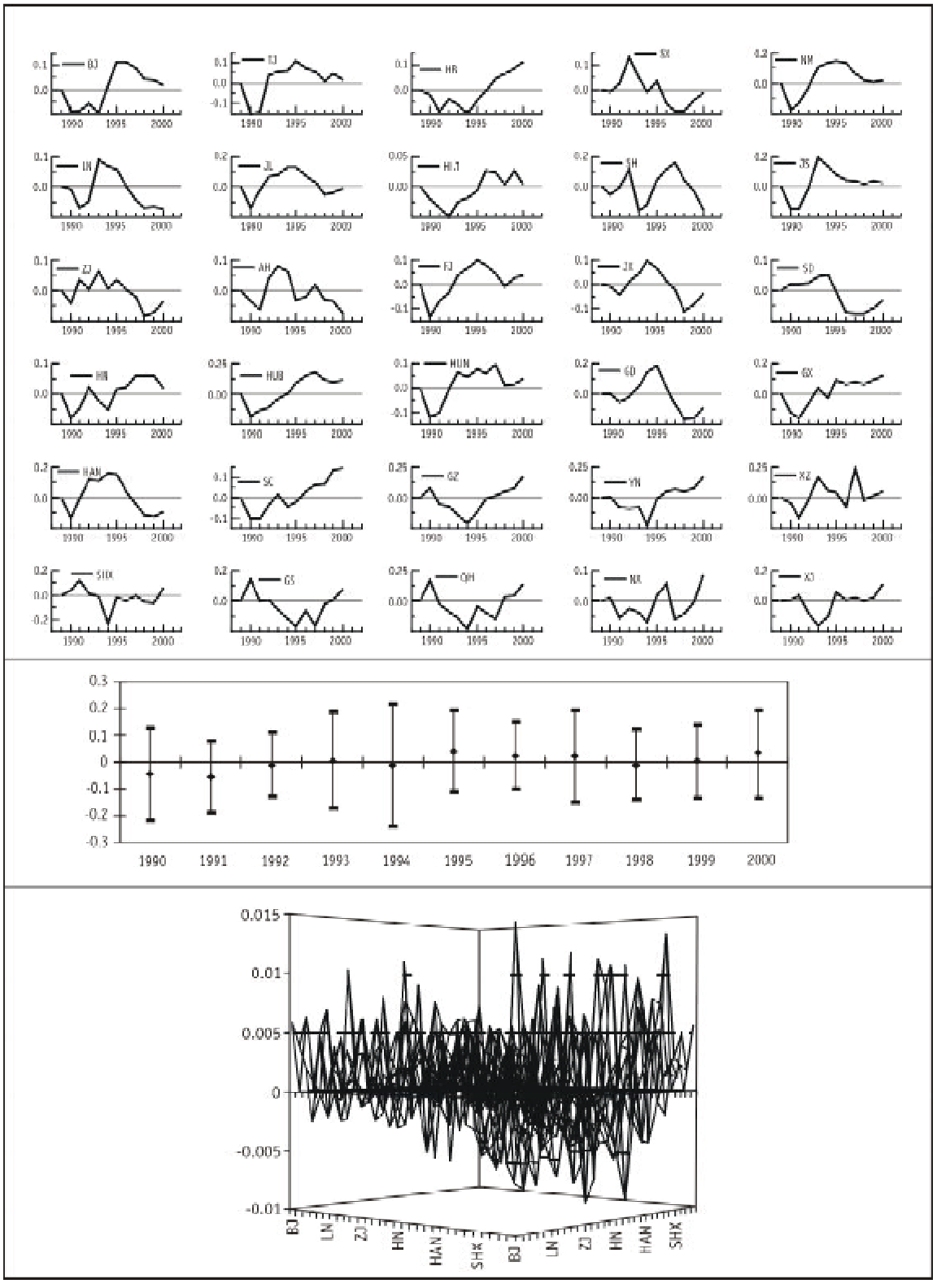


**Table 6: Rank Correlation Coefficients of Efficiency Measures (Standard Deviation: 0.1857)**

| 1st-Order Autocorrelation | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Year | (90,91) | (91,92) | (92,93) | (93,94) | (94,95) | (95,96) | (96,97) | (97,98) | (98,99) | (99,2000) |
| $z^{\tau}$ | 0.0425 | 0.1399 | 0.1199 | 0.0189 | -.1591 | 0.1693 | 0.3001 | 0.4478 | 0.4087 | 0.5453 |
| $z^{m}$ | 0.5835 | 0.3771 | 0.4656 | 0.8162 | 0.7219 | 0.6974 | 0.5907 | 0.6801 | 0.9350 | 0.7602 |
| Between $z^{\tau}$ and $z^{m}$ | | | | | | | | | | |
| 1990 | 1991 | 1992 | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 |
| -.00645 | -.1484 | 0.1858 | 0.180 | 0.0963 | -.2156 | 0.0345 | -.0145 | 0.1026 | -.1057 | -.0643 |
| Between $z^{m}$ and $\Lambda_i$ | | | | | | | | | | |
| Full sample: -0.08343 | | | | | Sub sample (92-2000): 0.04605 | | | | | |

Note: All rank correlation coefficients use Spearman's formula. In the calculation of rank correlation coefficients between $z^{m}$ and $\Lambda_i$, we take provincial means of $z^{m}$ for the appropriate sample size first before ranking them. Sample size is 30.

where $\bar{\delta}$ denotes sample mean and the index is normalized by max{$\Lambda$i}. The second line of (17) is based partly on the observation that depreciation rate data show little difference across province and time, and partly on the consideration that officially defined depreciation rates can be markedly different from the effective depreciation rates required by theoretical models like (11). As for $\alpha_i$, we choose to use the estimates from model (10a) rather than (15b) for the reason that the estimated $\alpha_i$ in (15b) are likely to contain certain heterogeneous effects due to the institutional factors.

Since we are unable to estimate $\sigma$ via the cost-minimization route because of the insignificance of the cost variable, we have to estimate $\alpha_i$ via a production function if $\sigma \neq 0$. Regarding the fact that most empirical studies of the PRC's aggregate production function assume $\sigma = 1$, we follow suit for simplicity and specify (8) by the Cobb-Douglas type of production function in a mixed panel form:

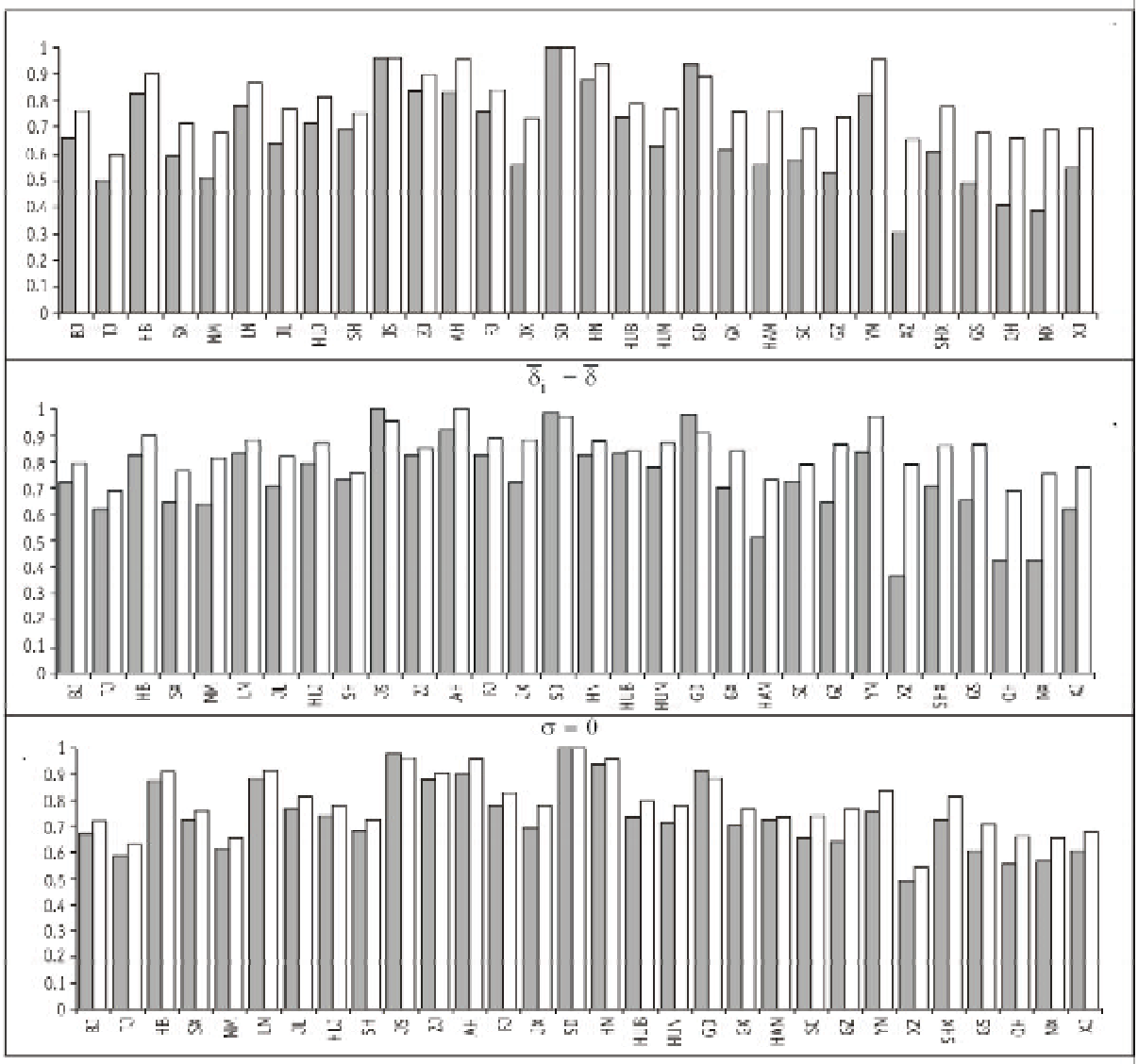
$$ln YI_{it} = \lambda_0 + \lambda_1 ln LI_{it} + \sum_{i=1}^{30} \alpha_i ln KI_{it} + u_{it} \qquad (8')$$

Due to lack of aggregate data on capital, we use data of the industrial sector here and assume that the spatial pattern of the estimated $\alpha_i$ applies to all the other sectors. Since the variables of (8') usually exhibit strong nonstationary property, $\alpha_i$ are taken as the long-run solution of a first-order ADL version of (8'), similarly specified as (10a'), and estimated by the combined GMM method. Two sets of $\alpha_i$ are calculated, one for sample 1988-1999 and the other for 1991-1999.[17] Correspondingly, two sets of $\Lambda_i$ are calculated under the three situations of (17) respectively, namely $\sigma$=1, $\bar{\delta}_i = \bar{\delta}$, and $\sigma$=0. The results are plotted in Figure 7, where the order of graphs goes with the line order of (17). We see from the figure that there is no considerable change in the general pattern of $\Lambda_i$ between different situations. The pattern appears to be in accord with what is usually perceived, namely coastal and southern provinces tend to be technologically more efficient than inland and western provinces. In particular, our result does not contradict Yao's estimates of

[17] We think it more appropriate to use the one-period lagged estimates of $a_i$ when calculating $\Lambda_i$ of (17). The two sets of estimated $a_i$ are from full sample of 1989-2000 and subsample 1992-2000, respectively.

technological inefficiency using micro firm data (2001). Moreover, provinces with relatively high PE tend to show better performance in their AE indices by and large. Rank correlation between $\Lambda_i$ and $z^m$ shows (see Table 6) that the two efficiency measures hardly relate to each other, verifying the postulate by Bai et al. (1997) that PE may not imply AE when firms' objectives are more complicated than profit maximization due to imperfect market environment.

**Figure 7: $\Lambda_i$ of (17):**




Legend: 1989-2000 shaded bars
1992-2000 nonshaded bars

## V. CONCLUSION

Excess investment at the aggregate level is a common phenomenon in CPEs. Has it disappeared during extensive economic reforms in many formerly CEPs? Evidence from the PRC suggests that the capital formation has outpaced GDP growth in the PRC over the last decade, especially during the periods of 1990-1994 and 1997-2000. Is this evidence of excess investment? In this study, we set up a model to estimate inefficiency in aggregate investment and to attribute the inefficiency to various economic and institutional factors. The model is used to analyze the PRC aggregate provincial-panel data.

Primarily, we have identified disequilibrium investment, i.e., deviations of actual investment from the market desired investment, to the ratio of investment to GDP, a result that confirms the previous findings by Sun (1998) and Song et al. (2001). The user cost of capital is found to play a negligible role, indicating that capital prices have not emitted market-clearing signals to reflect the true costs of investment demand. This finding is consistent with Stigilitz's observation (1996, 97) that, unlike in a pure market economy, firm managers in a transition economy tend to undertake grandiose investment projects, because their decisions generally do not bear the risks or costs of mistakes that they might make, but may, however, get credit for any achievement under their direction. A major factor sustaining this kind of behavior is incomplete and ambiguous property rights, which still prevails in the PRC firms.

Noticeably, our model design enables us to uncover how much disequilibrium investment can be explained by nonmarket-equilibrating institutional variables, which act as proxies of soft budget constraints. Fiscal deficit is found to contribute significantly to excess investment demand. In particular, provincial fiscal deficit appears to explain a slight and gradual declining return to scales observed from aggregate data. A network effect is also found to exacerbate disequilibrium investment, suggesting that provinces will not curb their investment desire until they join ranks with their regional leaders of excess investment. Both findings are consistent with the "federalism" argument by Huang (1996) and Qian and Roland (1998), as well as with the evidence previously presented by Zhang and Zou (1996) and Young (2000).

The essential advantage of our modelling approach is embodied in three clearly defined and estimable measures of investment efficiency. These measures enable us to draw distinction not only between inefficiency caused by resource misallocation and by underutilization of capital assets in production, but also between allocative inefficiency caused by imperfect market system and by firms' nonoptimal investment decisions. Estimates of the two AE measures suggest that severe underutilization of investment recourses is closely associated with governments' major attempts to stimulate aggregate demand and boost economic growth, and that there have been signs of improvement in firms' AE as the reforms deepen. These results support the views that decentralization and federalization generates mixed welfare effects (see Huang 1996, and Qian and Roland 1998). The third measure on PE is broadly in line with the pattern of regional development, with southern and coastal provinces more efficient than western provinces at large.

We must acknowledge that our efficiency measures have limitations. For example, the standard efficiency criterion that these measures are based on does not take into account the possibly positive externality of government nonprofit-seeking investment demand. Moreover, the measures have not explicitly allowed for the role of future expectation. In other words, our AE measures may not be able to indicate whether inefficient investment allocation at present will eventually become efficient

in the sense that it enhances the PRC's potential for future development. However, this kind of dynamic effect should not significantly bias our general results, as our measures are built upon disequilibrium investment from its long-run path. Efficiency is a normative concept after all. Model-based definitions and estimable measures should at least help to clarify previously confused views and disorganized evidence, and hopefully to reduce the gap between theoretical and empirical studies on the welfare implications of institutional changes in transitional economies.

## APPENDIX

**Main data sources:**

National Bureau of Statistics: *Statistical Yearbook of China* (SYC), *Industrial Economic Statistical Yearbook of China* (IESYC), *Statistics on Investment in Fixed Assets of China* (SIFAC), *Provincial Statistical Yearbook* (PSY), various issues.

China Finance Ministry: *Financial Yearbook of China* (FYC), various issues.

People's Bank of China: *Almanac of China's Finance and Banking* (ACFB), various issues.

**Variable definition and source:**

$I$: Fixed-asset investment at provincial level, SYC and SIFAC, adjusted to constant price by $P_I$

$Y$: GDP at provincial level, SYC, adjusted to constant price by $P_Y$

$P_I$: Price index of fixed-asset investment at provincial level, SYC

$P_Y$: Price index of GDP at provincial level, SYC

$r$: Real interest rate calculated by 3-5 year loan rates net of the growth rate of $P_I$ of one-year lag (proxy for expected inflation of investment goods), SYC and ACFB

$\delta$: Depreciation rate of fixed assets of state-owned industrial firms at provincial level, FYC and PSY (data for 1999 and 2000 unavailable, calculated using previous observations together with data of the net gross asset values of state-owned industries at provincial level from IESYC)

$\pi$: Tax is derived from total pre-tax profits minus total after-tax profits of industrial firms with independent accounting systems at provincial level, tax rate is then calculated using tax divided by value-added of the firms, SYC

$x_1$: Logarithm of net government debt, i.e., total government debt incurred minus total retirement of debt and interest payments (the net debt amounts approximately to the total government deficit); a series of central government deficit is also calculated, SYC

$x_2$: Logarithm of the ratio of provincial government expenditure to revenue, SYC

$x_3$: One-period lagged provincial Ii/Yi minus its regional average I/Y, standardized by the national average of I/Y

$x_4$: One-period lagged provincial per capita GDP minus its regional per capita GDP, standardized by the national per capital GDP, SYC, and PSY

*YI*: Value-added of Industry at provincial level, IESYC, 1989-1999

*LI*: Average employment of Industry at provincial level, IESYC, 1989-1999

*KI*: Net fixed assets of Industry at provincial level, IESYC, 1989-1999

**Abbreviation of provinces by region:**

| **Coastal region** | | **Central region** | | **Western region** | |
|---|---|---|---|---|---|
| BJ | Beijing | SX | Shanxi | SC | Sichuan |
| TJ | Tianjin | NM | Inner Mongolia | GZ | Guizhou |
| HB | Hebei | JL | Jilin | YN | Yunnan |
| LN | Liaoning | HLJ | Heilongjiang | XZ | Tibet |
| SH | Shanghai | AH | Anhui | SHX | Shaanxi |
| JS | Jiangsu | JX | Jiangxi | GS | Gansu |
| ZJ | Zhejiang | HN | Henan | QH | Qinghai |

FJ Fujian
SD Shandong
GD Guangdong
GX Guangxi
HAN Hainan
HUB Hubei
HUN Hunan
NX Ningxia
XJ Xinjiang

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