

Does fiscal decentralization promote economic growth? An empirical approach to the study of China and India

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Abstract

Purpose – The authors test the effect of expenditure decentralization and fiscal equalization on short- and long-run economic growth and estimate two-step generalized method of moment (GMM) simultaneous equations models, using panel data for China and India for the period 1985 to 2005. The authors estimate two simultaneous equations: a growth equation and equalization equation and find that expenditure decentralization has a negative and statistically significant effect at conventional levels on short-run economic growth for both China and India. However, the authors also find that this result is sensitive to the set of included explanatory variables. This leads the authors to conclude that expenditure decentralization has no effect on short-run economic growth for either country. The authors also find that expenditure decentralization has a positive and statistically significant effect on fiscal equalization for both countries but find no evidence that fiscal equalization affects short-run economic growth for either China or India. In contrast, the authors find that expenditure decentralization has a positive effect on long-run economic growth in the case of India, but not in the case of China. Finally, the authors report evidence that fiscal equalization has no effect on long-run economic growth in the case of China; however, the authors find that equalization has a positive and statistically significant at conventional levels effect on long-run economic growth in India.

Design/methodology/approach – The authors estimate two-step GMM simultaneous equations models, using panel data for China and India for the period 1985 to 2005. To examine the effect of fiscal decentralization (FD) policies on economic growth in China and India, the authors estimate two equations: a growth equation and an equalization equation. For the growth equation, the authors adopt a production-function-based model that is widely used in the empirical literature on growth; however, the authors do make some compromises with this specification due to the unavailability of certain data. For the equalization equation, the authors include variables that economic theory and empirical evidence suggest influence fiscal disparities among subnational governments which in turn influence the demand for horizontal fiscal equalization (HFE). To the extent possible, the authors employ the same econometric specification, variable constructions and sample periods for both China and India. The authors believe this strategy provides a more rigorous test of the FD hypothesis.

Findings – The authors find that expenditure decentralization has a negative and statistically significant effect at conventional levels on short-run economic growth for both China and India. However, the authors also find that this result is sensitive to the set of included explanatory variables. This leads to conclude that expenditure decentralization has no effect on short-run economic growth for either country. The authors also find that expenditure decentralization has a positive and statistically significant effect on fiscal equalization for



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Research limitations/implications – Due to the importance of FD policies, especially to many developing countries that are currently pursuing decentralization reforms, future research should examine the effect of FD on economic growth for other countries. Furthermore, although it would be difficult to do so, future research should examine whether FD promotes political stability on ethnically diverse countries.

Originality/value – To the best of the authors' knowledge, no one has examined the effect of FD policies on India's growth experience. What is more is that this is also the first of its kind to have a comprehensive empirical investigation into these two major developing countries with very interesting similarities and differences in FD policies. It is thus of great importance to examine the effect of expenditure decentralization and HFE on economic growth in China and India.

Keywords Fiscal decentralization, Economic growth, Horizontal fiscal equalization

Paper type Research paper

Introduction

China and India's intergovernmental fiscal systems (IGFSs) directly affect the well-being of over one-third of the world's population and are thus well worth studying. In addition to market reforms and trade liberalization, both countries are pursuing fiscal decentralization (FD) reforms [1]. Tiebout (1956) and Oates (1999) contend that FD can increase the allocative efficiency of the public sector thus potentially promoting economic growth [2]. In contrast, Prud'homme (1995) and Tanzi (1995) caution that there are potential risks associated with FD, including fiscal disparities among subnational governments and macroeconomic instability; either one or both of which may adversely affect a country's economic growth rate.

An important element of FD is expenditure decentralization whereby public services like primary education and healthcare, public safety and solid waste management to name just a few, are provided by subnational governments. In addition to expenditure decentralization, intergovernmental grants are often a major source of revenue for subnational governments in China and India [3]. The reason being that subnational governments in developing countries frequently lack the necessary tax base to self-finance their assigned expenditure responsibilities, particularly rural local governments, where subsistence agriculture is often the dominant economic activity. In such cases, the central government may use intergovernmental grants to address the fiscal gap between a subnational government's expenditure responsibilities and its ability to raise own tax revenues.

In addition to expenditure decentralization, China and India use intergovernmental grants to equalize subnational expenditures per capita or horizontal fiscal equalization (HFE). Since these transfers are financed with distortionary taxes, HFE may have an adverse effect on economic growth. Furthermore, equalizing grants themselves may be distortionary, thus adversely affecting economic growth. Therefore, we examine the effect of both expenditure decentralization and HFE on economic growth in China and India.

As discussed in greater detail below, the design and practice of FD policies in China and India involve compromises with the policy prescriptions of the normative literature on FD. These compromises may undercut the potential benefits of FD on allocative efficiency in the public sector and, in turn, adversely affect economic growth. Given the complexity of fiscal systems, the effect of FD policies on economic growth cannot be resolved by theory alone. The net effect of the countervailing forces of FD on economic growth is ultimately an empirical question.

To the best of our knowledge, no one has examined the effect of FD policies on India's growth experience, and the empirical evidence for the case of China is mixed. The conflicting results for China may reflect differences in sample periods, in econometric specifications and in construction of the regressors used in these studies [4]. To address this concern, we estimate identical econometric specifications using identical sample periods (1985–2005) and variable constructions. We estimate a short-run growth model, using two-step generalized method of moment simultaneous equations model (GMM SEM). We estimate a system consisting of two equations: a growth equation and an equalization (HFE) equation. We further assume that economic growth influences HFE and vice versa. Arguably, FD policies have a long-run, rather than a short-run, effect on economic growth and HFE. Therefore, we also estimate long-run models for China and India, using a two-step GMM five-year moving-average estimator. As with the short-run model, we assume that expenditure decentralization influences economic growth and HFE in two separate equations. As stated above, economic growth also influences HFE and vice versa. Therefore, in the growth equation, FD and HFE appear as among explanatory variables while in the equalization equation, FD and economic growth appear as among explanatory variables, both of which constitute the simultaneous equations model (SEM). We believe that estimating both short- and long-run models with identical econometric specifications, sample periods and variable constructions for two prominent decentralizing countries provides a more stringent test of the FD hypothesis than estimating a short-run model for a single country.

One of the greatest challenges in estimating this system of equations is devising a convincing identification strategy to address the potential endogeneity of expenditure decentralization, economic growth and HFE. The FD reforms in both China and India in the mid-90s create a plausible source of exogenous variation in both expenditure decentralization and HFE, which are the focus of this study. We also adopt an instrumental variables approach to identify the estimated coefficients of the potentially endogenous variables in our growth and equalization equations.

In the case of the short-run models, we find that expenditure decentralization has a negative and statistically significant effect at conventional levels on economic growth for both China and India. We also find that this result is sensitive to the set of included explanatory variables, which leads us to conclude that expenditure decentralization, at least within the range of expenditure decentralization observed in our samples, has no effect on short-run economic growth for either country. This leads us to reject that FD hypothesis that expenditure decentralization has a positive effect on short-run economic growth. We also find that expenditure decentralization has a positive and statistically significant effect on HFE in the case of both countries. However, HFE has no effect on short-run economic growth in either case. Finally, in the case of China, we find no evidence that expenditure decentralization influences long-run economic growth. In contrast, in the long-run model for the case of India, the combined direct and indirect marginal effects of FD on economic growth is positive for all values of FD. In other words, we cannot reject the FD hypothesis in the case of the long run for India. Finally, we find that HFE has a positive and statistically significant effect on long-run economic growth in the case of India, but no effect on China's long-run growth rate. These are interesting and important findings.

The remainder of this paper is organized as follows. In [Section 2](#), we compare China and India's IGFRs. [Section 3](#) summarizes the existing literature on the effect of FD on economic growth. [Section 4](#) describes our empirical strategy, sample period and variable constructions. [Section 5](#) discusses our empirical results, and [Section 6](#) concludes.

China and India's intergovernmental fiscal systems

There are interesting similarities and differences in the respective fiscal systems of China and India. We proceed below by briefly comparing the evolution of China and India's IGFSSs. For more elaborate descriptions of their fiscal systems, see [Rao \(2003\)](#), [Martinez-Vazquez and Rider \(2006\)](#), [Singh \(2009\)](#) and [Jin *et al.* \(2011\)](#) [5].

China's intergovernmental fiscal system

The People's Republic of China has a unitary form of government with five levels of government hierarchically arranged in a pyramid-like fashion with the central government at the apex of the pyramid [6]. FD has been a gradual process in China. In 1978, market and fiscal reforms started with the devolution of control over resources and decision-making power to subnational governments, to state-owned enterprises, to private enterprises and to individuals. Interestingly, China's fiscal system largely skips over the provincial level of government in favor of local government autonomy. [Zhang and Zou \(1998\)](#) contend that this fragmentation of authority among many local governments may prevent them from undertaking large infrastructure projects that confer benefits to a large region comprising many local governments. The possible lack of investment in infrastructure projects with significant regional spillover benefits may impede economic growth.

China's initial fiscal reforms led to uncontrolled decentralization and case-by-case bargaining between the central and local governments regarding revenue retention leading to sharp declines in total national revenues as a share of GDP and declines in the central government's share of total revenues. To address these issues, China initiated the tax sharing system (TSS) reforms in 1994. As a result of these reforms, the central government now takes the major share of tax revenues. Taxes exclusively assigned to local governments are low-yielding taxes, such as slaughter taxes.

China's IGFS is characterized by highly decentralized expenditures, highly centralized revenues, a high degree of subnational government transfer dependence and hidden and perhaps extra-legal subnational borrowing. Even though China's subnational governments are prohibited from borrowing, except upon special approval by the central government, almost all of them are circumventing these restrictions by borrowing "off the books" [7]. Ironically, uncontrolled spending by subnational governments due to soft-budget constraints may promote short-run economic growth [8]. If, however, the central bank is called upon to bailout insolvent subnational governments through monetary emissions, soft-budget constraints may result in macroeconomic instability which would have a negative effect on economic growth.

Another anomaly in China's fiscal system, according to the normative literature, is extra-budgetary finance by subnational governments. Extra-budgetary finance constitutes a fiscal dual track: one that is authorized by the legal framework of China's fiscal system and one that is not. The size of extra-budgetary funds grew rapidly in the 1980s and 1990s, finally becoming equivalent in size to budgetary funds in 1991 [9]. As a result of several policy reforms that shifted extra-budgetary funds to budgetary funds, the former has started to decline [10]. In 2005, extra-budgetary expenditures declined to about 16% as an average share of budgetary expenditures. Since off-budget funds are levied on the same tax base as budget funds, off-budget funds, like extra-budgetary revenues, are positively correlated with own-source revenues which are important sources of interregional disparities in per capita provincial expenditures in China [11].

In sum, transfer dependence, lack of genuine revenue autonomy and soft-budget constraints undermine fiscal discipline among subnational governments which may have detrimental effects on short- and/or long-run economic growth. And, as previously noted, by skipping over the provincial level of government in favor of local governments, China's expenditure assignments may be too decentralized to capture potential regional spillover

benefits from large-scale public infrastructure investments which may have a negative effect on economic growth.

India's intergovernmental fiscal system

India is a federal republic comprising one central government (the Union), twenty-eight states and seven union territories, including the National Capital Territory of Delhi. In contrast to China's IGFS, which skips over the provincial level of government in favor of local government autonomy, FD in India's fiscal system is "stuck" at the state level, despite the enactment of the 73rd amendment to the Constitution in 1992 which grants statutory authority to local bodies or *panchayats*. Nonetheless, the 73rd amendment does not grant *panchayats* with genuine expenditure autonomy. Rather, India's "decentralization" of expenditure assignments to the *panchayats* are concurrent with those of the states. Concurrent expenditure assignments, as opposed to clear and exclusive expenditure assignments which is prescribed by the normative literature, risk creating confusion about which level of government is responsible for delivering a given service. This makes it difficult for citizen-voters to hold public officials accountable. Given the size of India's states – the populations of the 14 largest states of India vary between 12.5 million residents and in excess of 200 million residents – it is arguable whether India is genuinely decentralized. The centralization of expenditure responsibilities at the state level may prevent India from reaping the benefits of "genuine" FD.

India's federal system has managed to maintain relatively peaceful relations in a highly diverse population. FD and democratic elections may be helping to contain potential interethnic and caste rivalries in India. The resulting political stability undoubtedly helps promote economic growth. This is no small achievement when one considers the difficulties many multi-ethnic states have in maintaining political stability and economic growth. In contrast, China's political stability is arguably a consequence of a relatively homogenous population – 90% of China's population is Han Chinese – and highly centralized political authority. Nonetheless, FD may also be contributing to China's political stability and economic growth, as well.

In any event, political stability undoubtedly helps promote economic growth however it is being achieved in these two countries. Unfortunately, testing this hypothesis is beyond the scope of this study because of the lack of a convincing counterfactual.

As previously noted, China's fiscal system is arguably too decentralized; whereas, India's system is arguably too centralized. Ironically, these contrasting features of China and India's fiscal systems may be having a detrimental effect on their respective economic growth rates. Regarding similarities in their fiscal systems, China and India's IGFSs are both characterized by insufficient revenue autonomy, transfer dependence and the accumulation of nontransparent debts which may be giving rise to soft-budget constraints. These shared features of China and India's fiscal systems conflict with the normative prescriptions of FD which, according to the theory, may have detrimental effects on their respective short-run and/or long-run economic growth rates.

Literature review

In this section, we review the empirical literature on the effect of FD on economic growth. To the best of our knowledge, there are no empirical studies of the effect of FD on India's economic growth rate, and the empirical literature on the effect of FD on China's economic growth rate provides conflicting evidence. Finally, there is limited empirical evidence on the influence of HFE on economic growth, a prominent exception being [Qiao et al. \(2008\)](#).

[Zhang and Zou \(1998\)](#) find that FD has a negative and statistically significant effect on China's short-run economic growth rate. Using data for the periods 1970–1993 and 1994–2002, respectively, [Lin and Liu \(2000\)](#) and [Ding \(2007\)](#) find that FD has a positive effect on China's short-run growth rate; however, these studies fail to account for the potential

endogeneity of FD in the growth equation. This may result in biased estimates. Finally, using a panel of provincial-level data for the period 1985 through 1998, [Qiao *et al.* \(2008\)](#) report evidence that China's FD reforms have a nonlinear effect on growth. For values of the index of FD between zero and 0.5, the joint marginal effect of FD and FD² on growth is positive, but the joint marginal effect is negative for values of FD greater than 0.5. Examining the cumulative distribution function of FD in the sample used by [Qiao *et al.* \(2008\)](#), we find that 90% of the observations in their sample have a value of FD greater than 0.5, meaning that FD has a negative effect on economic growth for all but 10% of the observations in their sample [12]. In an important contribution to the FD literature, [Qiao *et al.* \(2008\)](#) also report evidence of a trade-off between equalization and short-run economic growth.

Using a panel data set of 46 countries for the period 1970 to 1989, [Davoodi and Zou \(1998\)](#) investigate the relationship between expenditure decentralization and short-run economic growth. In contrast to the predictions of the FD hypothesis, [Davoodi and Zou \(1998\)](#) find a negative and statistically significant relationship between expenditure decentralization and short-run growth for developing countries in their sample. For developed countries, however, they find that expenditure decentralization has no effect on short-run economic growth. In other words, they reject the FD hypothesis for developing and developed countries alike.

In another study of the effect of expenditure decentralization on short-run economic growth, [Xie *et al.* \(1999\)](#) derive a simple model of endogenous growth with spending by different levels of government. They demonstrate how FD affects the long-run growth rate of an economy. They apply their theoretical model to the United States economy and report evidence that the existing spending shares for state and local governments are consistent with growth maximization. In contrast to the previous studies, they fail to reject the FD hypothesis in the case of the United States which has a relatively mature federal system.

We extend the existing literature on the effect of FD policies – expenditure decentralization and HFE – on economic growth in several ways. First, ours is the only empirical study of the effect of FD policies on India's economic growth rate. This fills an important gap in the literature. Second, we estimate our models for two prominent developing countries using identical sample periods, econometric specifications and variable constructions which facilitates the comparison of results between these two countries. We believe that this provides a more rigorous test of the FD hypothesis than previous studies. We also believe that a comparative study of China and India is interesting in and of itself. Third, the sample period of our data straddles the FD reforms of the early 1990s in both countries. These fiscal reforms provide a source of plausibly exogenous variation in FD and HFE which should help identify the estimated coefficients of the model. Fourth, in the case of China, we use the most complete data available on off-budget and extra-budgetary funds, which account for over one-half of local government revenues prior to the 1994 TSS reforms. Some of the previous studies either omit or use incomplete data on off-budget and extra-budgetary funds. This omission may result in biased estimates of the coefficients due to measurement error in a right-hand-side variable. Finally, ours is the first study to examine the effect of FD policies on China and India's long-run growth rate. Previous studies focus on the impact of FD policies on short-run economic growth. Again, we believe that our estimation of a long-run model fills an important gap in the literature.

We proceed below with a description of our empirical strategy.

Empirical strategy

To examine the effect of FD policies on economic growth in China and India, we estimate a SEM: a growth equation and an equalization equation. For the growth equation, we adopt a production-function-based model that is widely used in the empirical literature on growth. For the equalization equation, we include variables that influence fiscal disparities among

subnational governments which in turn influence the demand for equalization. To facilitate comparisons, we employ identical econometric specifications, variable constructions and sample periods for both China and India.

We estimate the following system of equations:

$$G_{it} = \alpha_0 + \alpha_1 E_{it} + \alpha_2 FD_{it} + \alpha_3 FD_{it}^2 + \alpha_4 K_{it} + \alpha_5 EMP_{it} + \alpha_6 GOVS_{it} + \alpha_7 CDEV_{it} + \alpha_8 FR_{it} + \mu_i + v_{it} \quad (1)$$

$$E_{it} = \beta_0 + \beta_1 G_{it} + \beta_2 FD_{it} + \beta_3 FD_{it}^2 + \beta_4 GOVS_{it} + \beta_5 CDEV_{it} + \beta_6 FR_{it} + \beta_7 SMW_{it} + \beta_8 FSR_{it} + \delta_i + \varepsilon_{it} \quad (2)$$

The subscript i indicates the province (state), and t indicates the year ($= 1985$ to 2005). G is the growth rate in real gross regional product (GRP) per capita, and E is the index of HFE, following the definition pioneered by [Qiao *et al.* \(2008\)](#). More specifically, E is calculated as follows. First, we compute the difference between a given province's (state's) per capita expenditures in year t and the sample mean of provincial (state) per capita expenditures in year t . By taking the absolute value of these figures, we convert them into distances from the mean. We normalize these figures by dividing by the average per capita provincial (state) expenditures in year t . Finally, we apply a minus sign to these figures in order to give the index of equalization an intuitive interpretation. That is as E approaches zero from below, the province's (state's) per capita expenditures are becoming closer to that of the average provincial (state) expenditures per capita.

The rate of FD (expenditure decentralization) is measured as the ratio of provincial (state) government expenditures per capita and central government expenditures per capita for China (India). In the case of China, provincial and central government expenditures include both budgetary and extra-budgetary expenditures. We use a quadratic specification of FD because related studies (e.g. [Qiao *et al.*, 2008](#)) report evidence that FD has a nonlinear effect on growth and HFE. Since FD may have a nonlinear effect on economic growth and equalization, it is important to account for this by estimating a quadratic specification.

As discussed in greater detail below, we find that the marginal effect of FD on short-run economic growth is similar for both countries. We believe that this consistency in results for two prominent decentralizing countries provides stronger support for our choice of econometric specification than if we had estimated the model for a single country.

The growth equations for China and India include a vector of control variables that are typically used in empirical growth models [\[13\]](#). Capital (K) is measured as the growth rate in total fixed assets. Labor (EMP) is measured by the growth rate in the labor force. According to Solow–Swan model of long-run economic growth, investment and the growth rate of the labor force have positive effects on the growth rate of real output. Provincial (state) government size ($GOVS$) is the share of provincial (state) budgetary and extra-budgetary expenditures in GRP. Public investment ($CDEV$) is the share of central government development expenditures in total expenditures. Economic theory and empirical evidence suggest that public infrastructure investment and government size, particularly expenditures on education and healthcare, have positive effects on economic growth.

Fiscal regime (FR), in the case of China (India), is a vector of time dummy variables set equal to 1.0 after 1994 (1991) when the fiscal reform for China (India) went into effect; and zero otherwise. Given the goals of these reforms, we expect this variable to have a positive effect on growth.

We also include a matrix of variables to control for natural disasters ($CATD_{it}$), which are expected to have an adverse effect on agricultural production. $CATD$ is set equal to 1.0 for the province (state) and year in which the catastrophe occurs and “0” otherwise. We also

include a time dummy variable set equal to “1” in 1989 which is the year of the Tiananmen Square incident (TIN) and zero otherwise; we also include a time dummy variable set equal to 1.0 in 1997 which is the year of the Asian financial crisis (ASIAC) and “0” otherwise. Both variables are included to account for the adverse economic shocks of these two events, and both are expected to have a negative effect on growth. In the interests of space, we do not report the estimated coefficients of these variables. They are available from the authors upon request.

In the equalization equation, we include the growth rate (G) and a quadratic specification of FD. As previously discussed, the effect of growth on equalization is ambiguous. On the one hand, FD is likely to give rise to interregional disparities in per capita income, which the growth literature shows has a negative effect on economic growth. On the other hand, subnational governments are heavily dependent on formula-based intergovernmental transfers that include equalizing factors in the formulas. These transfers should have a positive effect on HFE. However, transfer dependence may have a negative effect on economic growth. As in the case of the growth equation, we use a quadratic specification of FD to account for potential nonlinear effects of FD on HFE. Government size (GOVS) is expected to have a negative effect on equalization because a larger government may crowd-out private investment which impedes economic growth, narrows the tax base and decreases tax revenues. CDEV are equalizing transfers and therefore should have a positive effect on equalization. The formula used to distribute development funds ensures that this variable is exogenous.

We also include the share of “missing women” in the total population (SMW) in the equalization equation [14]. Using the method of Coale and Banister (1994), we calculate the number of hypothesized missing women for each province (state) and in each year in our sample. SMW is expected to have an adverse effect on equalization. For example, increased pensions are required for elderly men who never marry because of the adverse sex ratios in China and India.

Meloche *et al.* (2004) contend that the rate of fiscal self-reliance (FSR) is an important determinant of horizontal fiscal balance among subnational governments. FSR is the share of a province’s (state’s) own revenue in provincial (state) revenue receipts. We also include the vector of fiscal reform dummy variables, as defined above, in the HFE equation. Given the goals of these reforms, we expect this variable to have a positive effect on equalization. Finally, CATD, TIN and ASIAC are expected to increase fiscal disparities. In the interests of space, we do not report the estimated coefficients of these variables. These estimates are available upon request from the authors.

The terms μ_i and δ_i in equations (1) and (2), respectively, are unobserved, time-invariant provincial (state) effects. The terms v_{it} and ε_{it} , in (1) and (2) respectively, are idiosyncratic shocks that are time-varying and represent unobserved factors that change over time. In this simultaneous equation model (SEM), a change in any disturbance term of equation (1) leads to a change in the potentially endogenous variable E that it directly determines. This, in turn, changes the other potentially endogenous variable, namely G . A similar logic applies to the effects of a change in any disturbance term in equation (2). As a result, G and E are potentially endogenous.

To identify the parameters of the SEM, we must impose exclusion restrictions equal to the number of equations. In the growth equation, we exclude the share of missing women and the rate of fiscal self-reliance. In the equalization equation, we exclude the growth rate of fixed asset investment and the growth rate of the labor force. The excluded variables in the growth equation serve as instruments for the potentially endogenous variable “ E ” in the growth equation and likewise, the excluded variables in the equalization equation serve as instruments for the potentially endogenous variable “ G ” in the equalization equation. We also believe that FD and FD^2 are potentially endogenous. We use lagged

values of FD and FD^2 and cross products of the lag values of FD and FD^2 and higher moments of the share of missing women, fiscal self-reliance and lagged values of equalization as instrumental variables FD and FD^2 . To be valid, instrumental variables must be correlated with the potentially endogenous regressors and uncorrelated with the error term. In the discussion of the empirical results below, we report the results of a variety of specification tests that support our contention that the instrumental variables are valid.

The data for China come from the University of Michigan's China Data Center and covers the period from 1985 through 2005. The data on extra-budgetary revenues, extra-budgetary expenditures and central transfers come from the [People's Republic of China, State Council \(1996\)](#) (retrieved on October 2008 from www.mof.gov.cn). The data for off-budgetary funds are only available starting in 1985; hence, the decision to begin the sample period of this study in 1985. Fiscal data on India's states beyond 2005 are no longer available from India's Central Bank. Therefore, we estimate our models using data for the period 1985 through 2005. The sample period (1985–2005) covers the main years of previous studies and straddles the major reforms to the FRs of China and India that occurred in the early to mid-1990s. Straddling these reforms provides a plausible source of variation in FD and HFE which is helpful in identifying the estimated parameters of the models. Although more data is always preferred to less, we cannot completely rule out the possibility that extending the sample period of this study beyond 2005 would change our results. However, we believe that our main conclusions are likely to stand even if we extend the sample period to include more recent years because there have not been major FD reforms in either country since 2005.

There is concern that China's government may be manipulating economic data for political reasons. Measurement errors can lead to biased estimates and in unknown directions. According to [Rawski \(2001\)](#), the intentional falsification of data, in terms of GRP statistics in particular, is common at all levels of government in China. In contrast, [Chow \(2006\)](#) concludes that China's official statistics are generally reliable and consistent with China's economy, although some data must be used with caution, as with data for any country. In particular, the extra-budgetary data from the Ministry of Finance are considered to be reliable.

Our China sample includes data on 31 provinces of mainland China, including Tibet, but does not include data for the island governments of Hong Kong, Macau and Taiwan. In 1997, the municipality of Chongqing separated from Sichuan Province, and in 1988 Hainan separated from Guangdong Province. Data for Chongqing and Hainan Provinces are available for the years after the bifurcations created these new provinces in 1997 and 1988, respectively. Consequently, there are 638 observations in our sample. We drop eight observations due to missing values, resulting in an unbalanced panel consisting of 630 observations.

The main source for India's state fiscal and general national accounts data is the Reserve Bank of India (retrieved in December 2009 from www.rbi.org.in). All-India data are from the Central Statistical Organization, and population data are from India's Office of the Registrar General. Data on natural catastrophes are from Natural Disaster Management, Ministry of Home Affairs. The following union territories of India are excluded from our sample: Andaman and Nicobar Islands, Chandigarh, Dadra and Nagar Haveli, Daman and Diu, Lakshadweep and Pondicherry. Thus, there are a total of 29 state governments, including the national capital territory of Delhi, in our sample. These 29 states include the states of Chhattisgarh, Jharkhand and Uttaranchal which were established as a result of the bifurcation of the states of Andhra Pradesh, Bihar and Uttar Pradesh, respectively, in 2000. Therefore, the data for Chhattisgarh, Jharkhand and Uttaranchal are available only for the six-year period beginning in 2000. The resulting unbalanced sample would consist of 694 observations. However, complete data for Mizoram are only available for the period

beginning in 2000; complete data for Arunachal Pradesh and Goa are only available for the period beginning in 1986; and complete data for Nagaland, Sikkim and Delhi are only available beginning in 1994. Consequently, there are 524 observations in our sample for India.

To calculate the growth rate in real GSDP per capita, we lose one year of data in both samples, and we also lose a year of data due to the two-step GMM procedure used to estimate these models. As a result, the China (India) sample consists of 599 (451) observations.

Summary statistics for the resulting samples are reported in Table 1. Comparing the summary statistics, we see that China's (India's) average growth rate in GRP (GSDP) per capita is 9.3% (5.4). Interestingly, in light of the discussion of the reliability of China's data, the sample standard deviation of India's growth rates is substantially larger than that of China. The degree of HFE in China and India are very similar, -0.46 versus -0.52 , respectively, with somewhat more equalization in China than in India. The growth rate in fixed asset investment is substantially greater in China than in India; this undoubtedly reflects China's reliance on manufacturing and India's reliance on services. The growth rate in labor forces is very similar for both countries.

Empirical results

Before discussing our regression results, we begin by taking an initial look at our data by examining the evolution of expenditure decentralization and HFE during the sample period of this study.

The evolution of fiscal decentralization in China and India

The conventional measure of FD is the ratio of subnational expenditures per capita as a share of consolidated (central government and subnational government) expenditures per capita.

Variable	China (1985–2005)		India (1985–2005)	
	Mean	Standard deviation	Mean	Standard deviation
Growth rate in real GRP per capita	9.31	4.14	5.42	12.36
Index of horizontal fiscal equalization	-0.46	0.53	-0.52	0.59
Index of fiscal decentralization	3.29	2.52	1.75	1.57
Index of fiscal decentralization-squared	17.19	35.58	5.51	15.91
Growth rate in fixed asset investment	20.24	17.00	1.93	1.21
Natural logarithm of population	17.19	0.89	16.59	1.60
Central government development spending	46.13	12.21	48.13	5.14
Government size	19.81	9.62	32.82	26.94
Rate of fiscal self-reliance	39.11	10.73	50.35	26.60
Share of missing women	3.88	0.67	3.03	0.94
Fiscal regime ^a	0.62	0.49	0.76	0.43
Number of observations	599	599	451	451

Notes(s): ^aFor China, the FR dummy variable = 1.0 for years ≥ 1994 , zero otherwise. For India, the FR dummy variable = 1.0 for years ≥ 1991 , zero otherwise

Data sources for China

1) China Data Center of the University of Michigan and China's Ministry of Finance

2) Flood data is retrieved in April 2009 from <http://www.chinawater.net.cn/flood>, level B and above with "1", otherwise, "0"; other catastrophe data was retrieved on April 2009 from <http://zzys.agri.gov.cn>

Data sources for India

1) Directorate of Economics and Statistics of respective State Governments and for All-India – Central Statistical Organization

2) Reserve Bank of India

3) Natural disaster data was retrieved on October 2009 from <http://www.ndmindia.nic.in>

Table 1.
Summary statistics

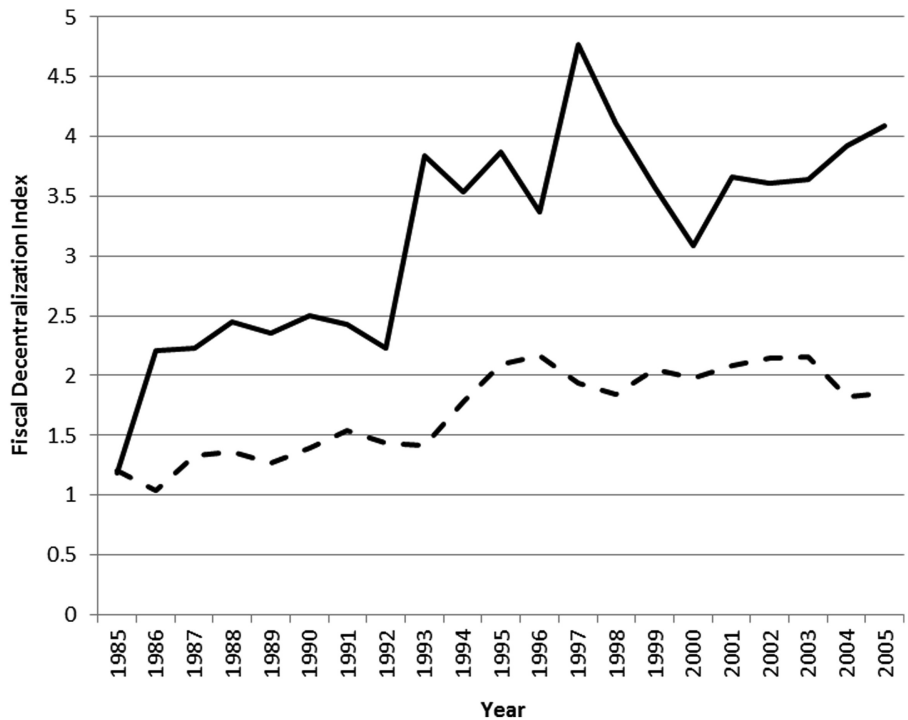
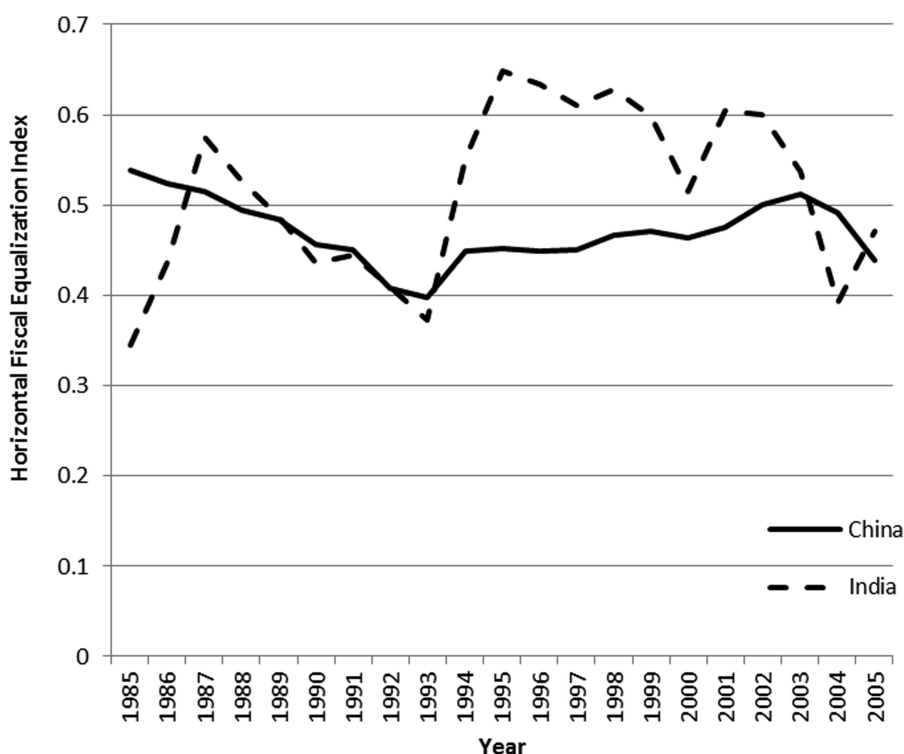


Figure 1.
Index of fiscal
decentralization for
China and India,
1985–2005

Key: China is indicated with a solid line, and India is indicated with a dashed line

Figure 1 shows the evolution of the average rate of expenditure decentralization among the provinces (states) of China (India) between 1985 and 2005. The solid line in Figure 1 represents the average rate of expenditure decentralization among the provinces of China for each year in our sample period, and the dashed line represents the average rate of expenditure decentralization among the states of India for each year in our sample period. Three patterns are evident in Figure 1. First, there is considerable variation over time in the rate of expenditure decentralization in both countries. Second, the average rate of expenditure decentralization is trending upward in both countries over time. More specifically, the average rate of China's expenditure decentralization varies between 1.0 and 4.0, meaning that the (unweighted) average of subnational expenditures per capita varies between 100% and 400% of central government expenditures per capita. Regarding India, the average rate of expenditure decentralization varies between 1.0 and 2.0, meaning that the (unweighted) average of subnational expenditures per capita varies between 100% and 200% of central government expenditures per capita. Third, there is considerably more variation over time in expenditure decentralization in China than in India during the sample period.

Figure 2 shows the evolution of the coefficient of variation(CV) of HFE among the provinces (states) of China (India) [15]. The solid line represents China's CV for each year in our sample period, and the dashed line represents the CV among the states of India. There is considerably more variation in the CV of the HFE index over time for India as compared to that of China.



The effect of
expenditure
decentralization

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Figure 2.
Index of equalization
(in absolute value) for
China and India,
1985–2005

Key: China is indicated with a solid line, and India is indicated with a dashed line

Regression results

We begin by discussing the estimates of our short-run growth models for China and India. This model examines the effect of contemporaneous values of FD and HFE on real GRP per capita; hence, our characterization of these models as short-run growth models. The estimated coefficients for the short-run growth models are reported in Table 2 [16].

For the case of China, the estimated coefficient of HFE (“ E ”) in the growth (“ G ”) equation is equal to 0.73 (standard error = 1.48); this estimate is statistically indistinguishable from zero at conventional levels of statistical significance. In the case of India, the estimated coefficient of HFE (“ E ”) in the growth equation (“ G ”) is -9.02 (S.E. = 14.61), which is also indistinguishable from zero at conventional levels of statistical significance. In other words, we find no evidence of a trade-off between HFE (“ E ”) and economic growth (“ G ”) for either China or India. The estimated coefficients of FD and FD^2 in the growth equation are negative and positive, respectively, and these estimates are statistically distinguishable from zero at conventional levels of significance for both countries [17].

According to the Sargan (1958) test of over-identifying restrictions and the Stock-Yogo (2005) weak-identification test, the instrumental variables used in the growth and equalization equations are valid instruments; that is, correlated with the potentially endogenous regressors and uncorrelated with the error terms. More specifically, the Sargan (1958) tests fail to reject the null hypothesis that the instrumental variables are exogenous. According to the Stock-Yogo (2005) weak-identification test, the relative instrumental

Table 2.
Two-step GMM
estimates of growth
and equalization for
China and India, 1985
to 2005

Variables	China (1985–2005)		India (1985–2005)	
	Growth equation	Equalization equation	Growth equation	Equalization equation
Growth rate in real GRP per capita	–	–0.003 (0.003)	–	0.002 (0.002)
Index of horizontal fiscal equalization	0.73 (1.48)	–	–9.02 (14.61)	–
Index of fiscal decentralization	–1.01* (0.59)	0.17*** (0.06)	–44.12*** (10.69)	0.24** (0.12)
Index of fiscal decentralization-squared	0.08*** (0.03)	–0.01*** (0.003)	2.21*** (0.83)	–0.03*** (0.009)
Growth rate of fixed asset investment	0.15*** (0.008)	–	9.80*** (2.04)	–
Growth rate of employment	–2.05 (3.42)	–	–45.66*** (9.03)	–
Government size	–0.03 (0.04)	–0.02*** (0.002)	0.09 (0.31)	–0.004 (0.003)
Central government development spending	0.02 (0.03)	–0.002* (0.001)	–0.29 (0.34)	0.007*** (0.002)
Share of missing women	–	–0.02 (0.02)	–	–0.01 (0.02)
Rate of fiscal self-reliance	–	0.002* (0.001)	–	–0.004** (0.002)
Fiscal regime (= 1 if year ≥ 1994; else 0)	3.23*** (0.87)	–0.16*** (0.05)	–	–
Fiscal regime (= 1 if year ≥ 1991; else 0)	–	–	–0.14 (3.61)	0.09*** (0.03)
Control variables for natural disasters	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Province (state) fixed effects	Yes	Yes	Yes	Yes
Number of observations	599	599	415	415
Weak-identification test	19.67 (9.37) ^a	13.47 (9.01) ^a	11.82 (7.77) ^a	5.14 (6.61) ^a
Sargan statistic	4.48	0.22	0.39	2.19
	<i>p</i> -value = 0.61	<i>p</i> -value = 0.64	<i>p</i> -value = 0.94	<i>p</i> -value = 0.34

Note(s): Robust standard errors are provided in parentheses

*** Statistically significant at the 1% level; ** at the 5% level; and * at the 10% level

The estimated coefficients of the catastrophe dummy variables are not reported here but are available from the authors upon request

^aThe critical value for the 10% maximal IV relative bias of the [Stock-Yogo \(2005\)](#) weak-identification test

variable bias is less than 10%, meaning that the instrumental variables are strongly correlated with the potentially endogenous regressors.

The evidence from the short-run models that FD has a nonlinear effect on economic growth is interesting and requires further interpretation. By differentiating (1) with respect to FD, we obtain the following expression for the marginal effect of FD and FD^2 on growth: $\delta G/\delta FD = \alpha_2 + 2\alpha_3 FD$. After substituting the estimated coefficients of α_2 and α_3 into the expression for a given country, we solve the resulting expression for the “critical” value, as it is often called, of FD that sets the expression above equal to zero. In other words, the critical value of FD equals $-\alpha_2/2\alpha_3$.

The critical values of FD, for China and India, respectively, are 6.31 and 10.02. Examining the cumulative distribution function of FD in our China (India) sample, we find that approximately 90% (98) of the observations have a value of FD less than the critical value. The joint marginal effect of FD and FD^2 on short-run economic growth is negative for all but 10% (2) of the observations in our China (India) sample. These findings for China and India lead us to reject the FD hypothesis. The policy implication of this finding is that China and India should increase expenditure decentralization, at least insofar as they are interested in using expenditure decentralization to promote short-run economic growth.

When interpreting our results from the SEM, it is important to be aware that FD and FD^2 potentially have direct and indirect effects on short-run economic growth. The direct effect of FD on growth, discussed above, is due to the influence of FD and FD^2 on economic growth in the growth equation. The indirect effect of FD on growth is due to the influence of FD and FD^2 in the equalization equation, which, in turn, has a potential effect on economic growth. As previously noted, HFE does not have a statistically significant effect at conventional levels on short-run economic growth for either country. Thus, in the short-run model, FD and FD^2 do not have an indirect effect on economic growth for either country. However, as discussed below, FD and FD^2 does have an indirect effect on growth in the long-run model for India.

As expected the growth rate of fixed asset investment has a positive and statistically significant effect at conventional levels on short-run economic growth for both countries. The effect of the growth rate of the labor force on short-run economic growth rate for China is negative but statistically indistinguishable from zero at conventional levels of significance. In the case of India, the estimated coefficient on the growth rate of the labor force is equal to -45.66 (S.E. = 9.03) which is statistically significantly different from zero at conventional levels. This finding may be surprising to some. However, there are long standing concerns about the potential adverse effects of India’s population growth rates; there are similar concerns for China, as well. Concern about “excessive” population growth rates are reflected in India’s substantial investments in family planning. Similarly, China’s former one-child policy likely reflected similar concerns about “excessive” population growth rates. Furthermore, it is theoretically possible for diminishing marginal returns of a variable factor, in this case labor and a fixed factor, say land, to make the marginal physical product of labor negative. This is a possible explanation for our finding that the growth rate in the labor force has a negative effect on the economic growth rate in the case of India [18].

Consistent with the goals of the China’s TSS reforms (labeled “fiscal regime” in Table 2), we find that they have a positive and statistically significant effect at conventional levels on economic growth. In contrast, the estimated coefficient of the 1991 reforms (labeled “fiscal regime” in Table 2) in India is statistically indistinguishable from zero at conventional levels of significance. This result is not altogether surprising in the case of India because, as previously discussed, the adoption of the 73rd amendment did not do very much to advance expenditure autonomy to *panchayats* (local governments) [19]. The remaining control variables in the growth equations for China and India are statistically indistinguishable from zero at conventional levels of significance.

Turning now to the equalization equations for China and India, we find that the effect of economic growth on HFE is statistically indistinguishable from zero at conventional levels of significance in the case of both countries. This is somewhat surprising because there is a widespread belief that the provinces of east China and the states of south India are leaving the provinces of west China and the states of north India behind in terms of growth in real GRP per capita.

The joint marginal effect of FD and FD^2 on equalization (HFE) is positive and statistically distinguishable from zero at conventional levels for both countries. In this case, the critical values for FD in the equalization equation is 8.5 and 4.0 for China and India, respectively, meaning that for values of FD less than the critical value, FD has a positive effect on equalization for 95% of the observations for both China and India. According to the Sargan test, we are unable to reject the null hypothesis that the instrumental variables for the potentially endogenous regressors in the equalization equation are uncorrelated with the error term, and according to the Stock-Yogo weak-identification test, the relative IV bias is less than 10%. In other words, we cannot reject the null hypothesis that the instrumental variables in the equalization equation are valid for both China and India.

In China, the estimated coefficients of government size and central government development spending in the equalization equation (HFE) are negative and statistically significant at conventional levels. The estimated coefficient of the rate of fiscal self-reliance in the equalization equation is positive and statistically significant effects at conventional levels. In the case of India, the effect of government size on equalization is statistically indistinguishable from zero at conventional levels. However, the effect of central government development spending on HFE is positive and statistically significant at conventional levels; whereas, the effect of the rate of fiscal self-reliance on HFE is negative and statistically significant at conventional levels. The effect on equalization of the share of missing women is statistically indistinguishable from zero at conventional levels of significance for both countries. Interestingly, the effect of the TSS reforms (labeled fiscal regime in Table 2) on HFE for China has a negative and statistically significant effect at conventional levels; whereas, the effect of the 1991 fiscal reforms on HFE in the case of India is positive and statistically significant at conventional levels.

Like Qiao *et al.* (2008), despite their stated conclusion to the contrary, we find that the direct marginal effect of FD on economic growth is negative for nearly every observation in our respective samples [20]. Thus, we conclude that there is very little evidence in support of the FD hypothesis when we account for the nonlinear effect of FD on short-run economic growth.

Following Leamer (1983) and Zhang and Zou (1998), we test the robustness of our results by conducting an extreme bounds analysis (EBA), which involves estimating the growth equation for every possible combination of the explanatory variables. We find that our results for the effects of expenditure decentralization on economic growth are fragile; that is, the signs and statistical significance of the estimated coefficients of FD and FD^2 are sensitive to the set of included explanatory variables. Zhang and Zou (1998) also conduct an EBA, and their findings are similar to our own. Like Zhang and Zou (1998), this finding leads us to conclude that FD has no short-run effect on economic growth for either China or India.

Thus far, we have followed the convention in the literature on FD and growth by assuming that FD has a contemporaneous effect on growth. Arguably, FD should have a long-run effect on economic growth as opposed to a short-run effect. After all, the Solow–Swan model is a long-run model of economic growth. Furthermore, to the extent that expenditure decentralization improves the quality of services like education and healthcare, these improvements may only manifest their impact on economic growth with a time lag, in other words, in the long run. To address this concern, we estimate a long-run growth model by estimating a five-year, moving-average specification of our two-step GMM SEM [21].

Variables	China (1985–2005)		India (1985–2005)	
	Growth equation	Equalization equation	Growth equation	Equalization equation
Growth rate in real GRP per capita	–	0.01*** (0.002)	–	0.01*** (0.002)
Index of fiscal equalization	1.08 (0.88)	–	12.00*** (4.24)	–
Index of fiscal decentralization	–1.21*** (0.39)	0.13*** (0.02)	–4.31 (3.14)	–0.15*** (0.05)
Index of fiscal decentralization-squared	0.08*** (0.02)	–0.01*** (0.001)	0.78*** (0.25)	–0.02*** (0.005)
Growth rate of fixed asset investment	0.17*** (0.01)	–	2.29** (1.03)	–
Growth rate of employment	–9.53*** (3.46)	–	34.70*** (5.63)	–
Government size	0.09** (0.04)	–0.01*** (0.002)	–0.26*** (0.07)	0.005*** (0.002)
Central government development spending	0.02 (0.04)	–0.001 (0.002)	0.81*** (0.22)	0.0004 (0.003)
Share of missing women in population	–	–0.01 (0.02)	–	–0.04*** (0.01)
Rate of fiscal self-reliance	–	–0.001 (0.001)	–	–0.002 (0.002)
Fiscal regime (=1 if year ≥ 1994; else 0)	5.22*** (0.96)	–0.08 (0.06)	–9.35*** (2.66)	0.33*** (0.05)
Fiscal regime (=1 if year ≥ 1991; else 0)	–	–	238	238
Number of observations	415	446	3.24	4.07
Sargan statistic	4.25	3.20	<i>p</i> -value = 0.20	<i>p</i> -value = 0.13

Note(s): Robust standard errors are provided in parentheses

*** Statistically significant at the 1% level; ** at the 5% level; and * at the 10% level

^aThe estimated coefficients of the catastrophe dummy variables are not reported but are available from the authors upon request

Table 3.
Two-step GMM
estimates of a five-year,
moving-average
specification for China
and India, 1985 to 2005

The estimated coefficients of our long-run growth model are reported in [Table 3](#) [22]. Focusing on the estimated coefficients of FD and FD² in the growth equation, the results are qualitatively similar to those obtained with the short-run growth model. More specifically, in the case of China, the estimated coefficients of FD and FD² in growth equation are negative and positive, respectively and statistically significantly different from zero at conventional levels. In the case of India, a joint test of the statistical significance of FD and FD² in the growth equation soundly rejects the null hypothesis of no effect at the 1% significance level.

Regarding the direct marginal effects of FD on economic growth, we find that expenditure decentralization (FD) has a negative effect on long-run economic growth for values of the index less than about 7.57 (2.77) for China (India). As previously noted, at least 90% of the observations in our two samples are less than these critical values, meaning that FD has a negative marginal effect on long-run economic growth for the vast majority of observations in our two samples.

Now, we examine the combined direct and indirect marginal effects of FD on long-run economic growth. As previously noted, the indirect marginal effect of FD works through its effect on equalization, which, in turn, has a potential influence on long-run economic growth. Since the estimated coefficient of HFE in the growth equation is statistically indistinguishable from zero at conventional levels in the case of China, it is not necessary to account for an indirect effect of FD on long-run economic growth. In the case of India, however, the effect of HFE in the growth equation is positive and statistically significantly different from zero at conventional levels; therefore, it is important to account for the indirect effect of FD on long-run economic growth.

Substituting (2) into (1) and differentiating the result by FD, we obtain the following expression:

$$\partial G/\partial \text{FD} = (1 - \alpha_1\beta_1)^{-1}\{\alpha_1\beta_2 + \alpha_1(1 - \alpha_1\beta_1) + 2[\alpha_1\beta_3 + \alpha_3(1 - \alpha_1\beta_1)]\text{FD}\}.$$

By plugging in the estimated coefficients for India, which are reported in [Table 3](#), into the expression above and evaluating the resulting expression for a variety of values of FD, we find that, in the case of India, the combined direct and indirect marginal effects of FD on economic growth is positive for all values of FD. In other words, we cannot reject the FD hypothesis in the case of the long run for India. This is an important insight into the influence of FD on economic growth.

Turning to the effects of the other variables in the long-run models, the signs, magnitudes and statistical significance of the estimated coefficients for China are similar to those obtained in the short-run model, with two notable exceptions. The estimated coefficient of the growth rate of employment is substantially larger in absolute value in the five-year, moving-average specification and statistically significant at conventional levels. Government size is positive and statistically significant at conventional levels in the five-year, moving-average specification. Finally, the estimated coefficient of the TSS reforms (labeled “fiscal regime” in [Table 3](#)) in China is somewhat larger in the five-year, moving-average specification compared to the estimate reported in [Table 2](#).

In the case of China’s long-run equalization equation, the results are very similar to the short-run results, except for central government development spending, the rate of fiscal self-reliance and the TSS reforms (labeled “fiscal regime” in [Table 3](#)) are statistically indistinguishable from zero at conventional levels of significance.

Turning to the other variables in India’s growth equation, there is a sign change in the case of the growth rate of employment in the five-year, moving-average specification compared to the estimate for the short-run model. The estimated coefficient of government size is negative and statistically significant at conventional levels, and the estimated coefficient of central

government development spending is positive and statistically significant at conventional levels.

In the long-run model, the estimated coefficient of the FR in the equalization equation for India is negative and statistically significant at conventional levels. The estimated coefficients in India's equalization equation are qualitatively similar to those reported for the short-run model, with only a few exceptions. Government size has a positive and statistically significant effect at conventional levels on equalization in the long-run model. Central government development spending and the rate of fiscal self-reliance are statistically indistinguishable from zero at conventional levels of significance, and the share of missing women is negative and statistically significant at conventional levels.

We also estimate a dynamic model of economic growth for China and India. This model uses a three-year lag of the dependent variable in each equation [23]. Focusing on the direct marginal effect of FD on economic growth, the results are very similar to those obtained for the short- and long-run models. After controlling for the one-, two- and three-year lags of the dependent variable, FD and FD^2 have the same sign pattern in the growth equation as in the previous specifications for both countries. While the estimates of these coefficients for China are statistically significant at conventional levels, in the case of India, they are statistically indistinguishable from zero at conventional levels. The lack of statistical significance of FD in the dynamic model for India may be due to the sample attrition resulting from the bifurcation of several states during the sample period. As before, the negative coefficient of FD dominates the positive coefficient of FD^2 in the growth equation for China, meaning that the joint effect of FD and FD^2 have a negative effect on economic growth in the dynamic model. As for the effect of lags of the growth rate, the one-year lag has the largest effect and is positive, the two-year and three-year lags are small and negative; all three lags are statistically distinguishable from zero at conventional levels of significance for both countries.

Conclusions

Using data for the period from 1985 to 2005, we estimate short- and long-run growth models to examine the effect of FD on economic growth and equalization for China and India. To facilitate the comparison of results for these two countries, we use identical sample periods, econometric specifications and variable constructions. We believe that this approach provides a rigorous test of the FD hypothesis.

In the case of the short-run growth model, we find that the direct marginal effect of FD on economic growth is negative (positive) for values of FD less than (greater than) 6.31 (10.02) for China (India). To examine the robustness of these results to the set of included explanatory variables in the short-run growth equation, we conduct an EBA. We find that the signs and statistical significance of FD and FD^2 in the growth equation are sensitive to the set of included explanatory variables. This finding cautions us to interpret our results with care. With this caveat in mind, the preponderance of the evidence appears to reject the FD hypothesis that expenditure decentralization has a short-run effect on economic growth in the case of both China and India. Interestingly, the policy implications of these estimates of the direct marginal effects of FD and FD^2 suggest that both China and India could increase their short-run growth rates by increasing expenditure decentralization.

Regarding the direct marginal effect of FD on economic growth in our short-run model, our findings are consistent with those of [Zhang and Zou \(1998\)](#) and, we contend, with those of [Qiao *et al.* \(2008\)](#), as well, despite their stated conclusion to the contrary. In contrast to [Qiao *et al.* \(2008\)](#) who report evidence of a trade-off between equalization and growth, we report evidence that equalization has no effect on short-run economic growth for both China and India. We also report evidence that equalization has no effect on long-run economic growth in

the case of China. However, we find that equalization has a positive and statistically significant at conventional levels effect on long-run economic growth in the case of India.

Why do the short-run models for China and India reject the FD hypothesis? It is arguable whether the FD hypothesis should apply in the short run. Furthermore, as previously discussed, the design and implementation of FD in China and India are not consistent with the normative principles of FD. Indeed, promoting allocative efficiency may not be the principal aim of FD in China and India. Rather, the guiding paradigm of FD in these two countries may be what [Noiset and Rider \(2011\)](#) refer to as “union preserving federalism”. In this view, limited revenue autonomy, transfer dependency and subnational borrowing, which characterize FD in China and India, increase the dependency of regions on the central government thereby limiting separatist tendencies in diverse countries with weak national identities. Union preserving federalism may explain why these two countries have opted for fiscal structures which in many respects violate the normative prescriptions of FD and result in our finding that FD has no effect on economic growth.

Due to the importance of FD policies, especially to many developing countries that are currently pursuing decentralization reforms, future research should examine the effect of FD on economic growth for other countries. Furthermore, although it would be difficult to do so, future research should examine whether FD policies as they are currently designed and implemented promotes political stability in ethnically diverse countries.

Notes

1. By fiscal decentralization we mean the devolution of specific government functions by the central government to subnational governments with the administrative authority and fiscal revenue to perform those functions.
2. In addition to promoting allocative efficiency in the public sector, [Qian and Weingast \(1997\)](#) contend that fiscal federalism can be market preserving which could further economic growth by increasing confidence in property rights. Our empirical strategy does not provide a test of this hypothesis.
3. According to the normative literature, there are four pillars of FD: expenditure assignments, revenue assignments, intergovernmental transfers and regulation of subnational borrowing. For a complete description of the four pillars of IGFS in India, see [Bahl et al. \(2005\)](#).
4. In China, the expenditure responsibilities and the assignment of revenue sources of local governments are specified by the central government. Off-budget and extra-budgetary funds are expenditures and revenues that are outside of the central government’s legal framework of FD. In the case of China, some of these studies either omit or use incomplete data on off-budget and extra-budgetary funds which account for over one-half of local government revenues. We use the most complete data available on off-budget and extra-budgetary funds.
5. [Rao \(2003\)](#), [Martinez-Vazquez and Rider \(2006\)](#), [Singh \(2009\)](#) and [Jin et al. \(2011\)](#) are descriptive studies of China and India’s fiscal systems; whereas, [Zhou and Zhang \(1998\)](#), [Lin and Liu \(2000\)](#), [Ding \(2007\)](#), [Qiao et al. \(2008\)](#) and the current study test the FD hypothesis.
6. See Figure 4.1 of [Qiao and Shah \(2006\)](#) for the structure of government in China and its political system.
7. [Qiao and Shah \(2006\)](#) describe the off-the-book channels of local government borrowings in China.
8. In the context of intergovernmental fiscal relations, a soft-budget constraint arises whenever the central government cannot commit to a no bailout policy for insolvent subnational governments. [Qian and Roland \(1998\)](#) describe the potential economic risks created by soft-budget constraints among subnational governments.
9. See [Wong \(1998\)](#) for the size, evolution and detailed categorization of off-budget funds.
10. See, for example, State Council Document No. 29 (1996).
11. See, for example, [Wong \(1998\)](#) for a detailed empirical analysis of extra-budgetary funds.

12. By not accounting for the nonlinear effect of FD on economic growth in the interpretation of their results, they mistakenly, in our opinion, conclude that FD has a positive effect on economic growth. This leads them to fail to reject the FD hypothesis. As the foregoing analysis shows the conclusion that FD has a positive effect on economic growth only applies to 10% of the observations in their sample. As discussed below, our results are very similar to those of [Qiao *et al.* \(2008\)](#). Where we appear to differ is in our interpretation of our respective results. We would like to thank Qiao, Martinez-Vazquez and Xu for sharing their data with us, which permitted us to make these calculations with their sample.
13. More specifically, the Solow–Swan model explains long-run economic growth as a function of net investment (i.e. the growth rate in capital accumulation net of the rate of depreciation), the growth in the labor force and the growth rate of technological change. This model was developed independently by [Solow \(1956\)](#) and [Swan \(1956\)](#). Due to the lack of data, we do not include a measure of the growth rate in technological change in the growth equation. However, we control for time fixed effects in the growth rate equation which should absorb the effect of long-run technological change on economic growth thus mitigating any bias due to an omitted variable.
14. Missing women is the shortfall in the number of women relative to the expected number of women in a region or country. See, for example, [Jiang *et al.* \(2005\)](#) for the adverse effects of missing women.
15. The coefficient of variation of HFE is equal to the variance of the index of equalization divided by its mean; the resulting quantity is then multiplied by 100.
16. We estimate our SEMs using the `xtivreg2` STATA command ([Schaffer \(2010\)](#)).
17. We also estimate the model using revenue decentralization as an alternative measure of FD. The results are similar to the results for expenditure decentralization in the growth equation for China and India in terms of both signs and statistical significance.
18. In fact, visitors to India are often struck by the apparent over staffing in many workplaces with large numbers of employees that do not seem to be very productive.
19. There are exceptions to this generalization about the effect of the 73rd amendment. For example, Kerala and West Bengal have granted their local governments with considerable fiscal autonomy.
20. [Qiao *et al.* \(2008\)](#), page 123) conclude the following: “[O]ur results show that, consistent with the logic of China’s fiscal reform, FD improved economic growth . . . a higher level of decentralization led to a higher growth, but as expected, this relationship was nonlinear.” Their main results, which are reported in Table 4 (2SLS fixed effect results for the sample period 1985 to 1998), the estimated coefficients of FD and FD2 are 367 (t -value = 3.23) and –368 (t -value = –3.85), respectively. Evaluating the expression above for the joint marginal effect of FD on economic growth and evaluating it at the mean value of FD (= 0.7) in their sample, we find that the joint marginal effect of FD is negative. At the very best, they report weak support for the FD hypothesis that expenditure decentralization has a positive effect on economic growth.
21. Although the effects of FD on growth may take place over even longer periods of time than five years, say ten- or twenty-year periods, we do not have a sufficiently long panel to test for such effects.
22. There is substantial attrition in the sample for India because it is not balanced due to the bifurcation of three states in the mid-1990s.
23. For the sake of brevity, we do not report the results from the dynamic model; however, they are available from the authors upon request.

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