Health Investment, Health Outcomes and Economic Growth in China: An Applied Macroeconomic Analysis

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Abstract

Health is regarded as an important dimension of human development and its crucial role as human capital has been long recognised; however, the strategic role of health investment remains undervalued in less developed regions. China is a very good example of this. Along with the economic transition in the past three decades, China's health care system has also undergone several major changes. A key policy change was a shift of the main responsibility of health care finance to the individual in the early 1990s and a reversal of this in the early 2000s after the SARS epidemic. Possibly as a result of this, the population's health status is reported to be performing below its potential. This thesis conducts a thorough empirical investigation of the various and varied relationships between health investment, health outcomes and economic growth in China, using publicly available macro data; thus it makes a significant contribution to the academic literature, as well as provides evidence for policy-makers in both China and other countries in transition.

More specifically, this thesis empirically studies (1) the long-run relationship between health investment and economic output; and (2) the short-run nexus between health expenditure behaviours, health status and economic fluctuations, through both national time series and provincial-level panel data from mainland China. Overall, the results suggest that health investment has a significant and positive role in explaining the level of economic output in the long-run; and an estimated stronger effect in inland China may further highlight diminishing returns in health investment. The short-run results also suggest a significantly pro-cyclical movement of government health expenditure with economic growth. Furthermore, child health outcomes and infectious disease incidence are also significantly associated with economic fluctuations.

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- The 32nd Australian Conference of Health Economists (September 2010, Sydney); Oral Presentation; Health Investment and Economic Output in Regional China.
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List of Abbreviations

2SLS	Two-stage Least Squares
3SLS	Three-stage Least Squares
ADF	Augmented Dickey-Fuller
AMG	Augmented Mean Group
AO	Additional Outlier
ASR	Adult Survival Rate
BN	Beveridge-Nelson
CBB	Continuous-Path Block Bootstrap
CCR	Canonical Cointegrating Regressions
CD	Cross-section Dependent
CDR	Crude Death Rate
CHNS	China Health and Nutrition Survey
CMS	Cooperative Medical Scheme
CPI	Consumer Price Index
CRS	Constant Returns to Scale
CVD	Cardiovascular Disease
DALYs	Disability-Adjusted Life Years
DOLS	Dynamic Ordinary Least Squares
ECT	Error-Correction Term
FDI	Foreign Direct Investment
FE	Fixed-Effects
FMLS	Fully Modified Least Squares
GDP	Gross Domestic Product
GH	Gregory and Hansen
GHE	Government Health Expenditure
GIS	Government Insurance Scheme
GLS	Generalised Least Squares
GMM	Generalised Method of Moments
GNP	Gross National Product
GRP	Gross Regional Product
GVA	Gross Value Added

HAC	Heteroskedasticity and Autocorrelation Consistent
HDI	Human Development Index
HLM	Hierarchical Linear Modeling
IMF	International Monetary Fund
IMR	Infant Mortality Rate
IO	Innovational Outlier
IV	Instrumental Variable
LIS	Labor Insurance Scheme
MDGs	Millennium Development Goals
MMR	Maternal Mortality Ratio
MRW	Mankiw-Romer-Weil
MSAs	Medical Savings Accounts
NL2SLS	Non-linear Two-stage Least Squares
NLS	Non-linear Least Squares
NMR	Neonatal Mortality Rate
NRCMS	New Rural Cooperative Medical Scheme
ODC	Overseas Development Council
OECD	Organisation for Economic Co-operation and Development
OLS	Ordinary Least Squares
OOP	Out-of-Pocket
PHE	Private Health Expenditure
PIM	Perpetual Inventory Method
PNM	Perinatal Mortality Rate
PP	Phillips and Perron
PPP	Purchasing Power Parity
PQLI	Physical Quality of Life Index
PV	Perron and Vogelsang
QS	Quadratic Spectral
RCDI	RUC China Development Index
SARS	Severe Acute Respiratory Syndrome
SUR	Seemingly Unrelated Regression
TB	Tuberculosis
TFP	Total Factor Productivity
U5MR	Under-Five Mortality Rate

UEBMI	Urban Employee Basic Medical Insurance
UN	United Nations
UNDP	United Nations Development Programme
URBMI	Urban Resident Basic Medical Insurance
UV	Ultraviolet
VAR	Vector Autoregression
VEC	Vector Error-Correction
WHO	World Health Organization

CHAPTER 1

INTRODUCTION

1.1. BACKGROUND AND MOTIVATION

Health has always been considered as a key dimension of development. In 1976 the Physical Quality of Life Index (PQLI) was developed by the Overseas Development Council (ODC) as a supplementary indicator of Gross National Product (GNP) per capita to measure the progress of capability to meet basic human needs (Overseas Development Council, 1977). The design of PQLI (which combines three indicators - life expectancy at age one, infant mortality and literacy - into a single index) reflects the recognition that GNP per capita alone cannot necessarily reflect widespread well-being (Grant, 1981). However, the PQLI is not without drawbacks. For example, Larson and Wilford (1979) shows that the three input variables are highly correlated, and any one of the three variables could be used to rank life quality.

The descendant of the PQLI is the most well-known well-being indicator, the Human Development Index (HDI), proposed by the United Nations Development Programme (UNDP) in its first annual Human Development Report in 1990. The HDI is composed of three dimensions[1]: a long and healthy life (life expectancy at birth), knowledge (adult literacy rate and gross enrolment ratio) and a decent standard of living (Gross Domestic Product (GDP) per capita) (United Nation Development Programme, 1990). In several relevant indexes (e.g. Inequality-adjusted Human Development Index, Gender-related Development Index, Human Poverty Index) developed by the UNDP, a long and healthy life is always a key component.

Although it is now widely accepted that good health is a basic capability and human right (Sen, 1989; Sen, 2008; Clark, 2008), the strategic role of health investment is still under-valued in less developed regions. As stated in the preface of the World Health Organization (WHO)'s Commission on Macroeconomics and Health (WHO, 2001), "Although health is widely understood to be both a central goal and an important outcome of development, the importance of investing in health to promote economic development and poverty reduction has been much less appreciated." Mr Chen Zhu, the Minister of Health of China, also pointed out "a traditional misconception is that spending on health is a social burden, instead of being a strategic investment essential for each nation's socioeconomic development" (Chen, 2010, page 1429).

[1] In the latest version published in the Human Development Report 2010, the knowledge/education dimension has been replaced by mean years of schooling and expected years of schooling; the living standards dimension has been replaced by GNP per capita. There are also country-specific HDI that use different variables with those three dimensions. For example, the American HDI considers educational degree attainment and school enrolment in the knowledge dimension, and uses median earnings to measure the standards of living (see www.measureofamerica.org/human-development for details).

The nexus between health and wealth of a nation has been found to be a two-way relationship (Bloom and Canning, 2000). The classical empirical studies focus on the causal link running from economic growth to health improvement, i.e. whether *the wealthier the healthier*. The majority of results find a positive effect of economic growth on health (Pritchett and Summers, 1996); furthermore, the effect is found to be influenced by the distribution of wealth (Biggs *et al*. 2010). On the other hand, with the recognition that health is a form of human capital (Schultz, 1961; Mushkin, 1962; Fuchs, 1966; Grossman, 1972), investment in health and thus an improvement in health status can also have a positive effect on economic growth, i.e. *the healthier the wealthier*[2].

While the above discussion highlights the health-growth relationship in a long-run perspective, the second area of interest focuses on the relationship from a short-run horizon, i.e. the nexus between economic fluctuations and health outcomes. Dating back to the research on economic fluctuations and suicide rates (Durkheim, 1952), much empirical research has been undertaken which studies the link between economic cycles and health outcomes (e.g. general age-adjusted, age-specific and disease-specific mortality rates). Recent economic crises have caught the attention of several studies, focusing on developing countries, where maternal and child health outcomes are mainly of interest (e.g. maternal mortality ratio, infant, neonatal and post-neonatal mortality rates). Although recent literature using data from developed countries has tended to suggest *recessions are good for health*, the results from developing countries' data present a different story, that *recessions are bad for health*. Different health spending behaviour is perhaps one reason for this. The relationship between economic fluctuations, health expenditure behaviour and health outcomes are inconclusive, depending on the chosen health outcomes and the region

[2] See Spence and Lewis (2009), Chapter 2 and 3 for a detailed summary of recent literature and new evidence.

of interest[3].

Research on the relationship between investment in health and economic growth can provide useful policy relevant evidence. Due to heterogeneity across countries, country specific analysis is especially relevant. Currently, the evidence focusing on the world's largest developing country, the People's Republic of China (hereafter China or mainland China), is still scarce.

This thesis aims to fill this gap providing new and novel analyses of long run and short run effects – firstly, using both national time-series and provincial-level panel datasets to empirically investigate the contribution of health investment on economic output in China in a relatively long-run perspective. Secondly, to examine short-run effects, this thesis analyses the cyclical behaviours of health investment during economic fluctuations and provides new evidence regarding the relationship between economic fluctuations and health outcomes in China.

In the rest of the Introduction, the trends in economic growth, the background of health care reform, and the progress in health status in China are briefly described.

1.1.1. Trends in National Income in China

The economic history of mainland China can be divided into four different stages, represented by shifts in economic policy: a centrally planned economy (1953-78); a planned commodity economy (1979-92); a socialist market economy (1993-99); and

[3] See Suhrcke and Stuckler (2012), Chapter 4 and 5 for a review and new evidence.

post-industrialism (2000-onwards) (Polsa *et al.*, 2006). Despite many political events, China achieved impressive economic growth during the pre-reform centrally planned period. Although there were marked fluctuations in growth over time, national income in real terms quadrupled from 1953 to 1978, with an annual growth rate of 5.68 per cent, 3.68 per cent higher than population growth (Yao, 1994). Following the economic reform of the so-called “open-door policy”, the Chinese economy has experienced even greater growth over the last three decades. Prior to 1979, China’s per capita GDP grew somewhat faster than India, but lagged dramatically behind several other countries, for example, Japan and the United States. A different picture emerged after the reform (see Table 1.1). The share of China’s GDP in the world has also increased from around 2.2% in 1980 to 12.9% in 2009, according to the World Economic Outlook Database (September 2011).

Table 1.1. China’s per capita GDP as a percent of figures for USA, Japan and India, 1913-2009.

	1913	1952	1980	1990	2000	2009
United States	9.8	2.7	2.1	3.4	6.7	15.0
Japan	37.6	11.8	3.0	4.2	9.4	21.1
India	77.6	42.6	59.9	90.2	155.1	218.9

Sources: 1913: Maddison (2002); 1952: Brandt and Rawski (2008); 1980-2009: World Economic Outlook Database (September 2011).
Notes: Calculation based on current international dollar, measured at purchasing-power-parity (PPP). Annual figure for each comparator nation equals 100.

However, along with this increased trend of wealth, substantial regional disparities exist. Mainland China consists of 22 provinces, 5 autonomous regions and 4 municipalities (hereafter, 31 provinces). The municipality of Chongqing separated from Sichuan Province in 1996. Similar to Fu (2004) and Keng (2006), Coastal China in this thesis refers to

Beijing, Tianjin, Hebei, Liaoning, Shanghai, Jiangsu, Zhejiang, Fujian, Shandong, Guangdong, Guangxi and Hainan (red part in Figure 1.1); all remaining provinces are classified as Inland China (green part in Figure 1.1).

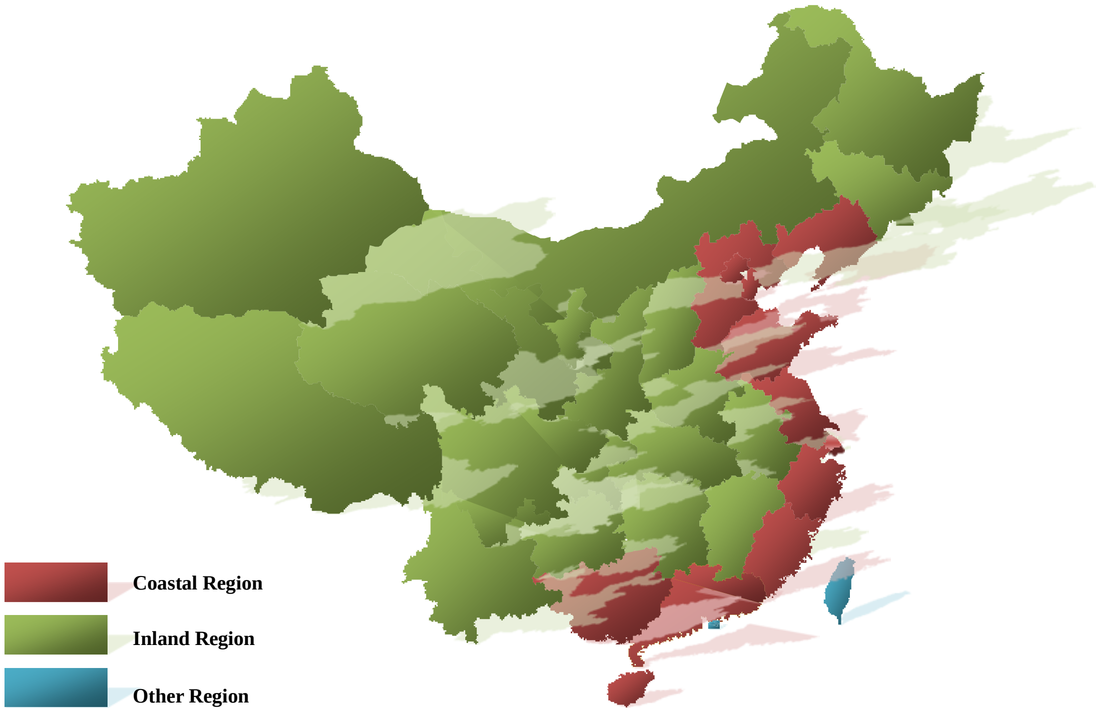

Figure 1.1. Map of China regions.

Using current values of data from 2009 for example, Figure 1.2 shows the share of three income indicators (i.e. GDP per capita, per capita disposable income of urban residents and per capita net income of rural residents) in each province as a percentage of the figures for Shanghai. Firstly, a clear intra-province disparity can be found: the GDP per capita of Guizhou province (the lowest in year 2009) is only 13.1% of the figures for Shanghai. Secondly, there is also a rural-urban disparity. The per capita net income of rural residents as a percentage of per capita disposable income of urban residents ranges from 23.4% (in Yuannan and Guizhou) to 43.64% (in Beijing). More detailed studies on the regional development in China can be found in Démurger *et al.* (2002).

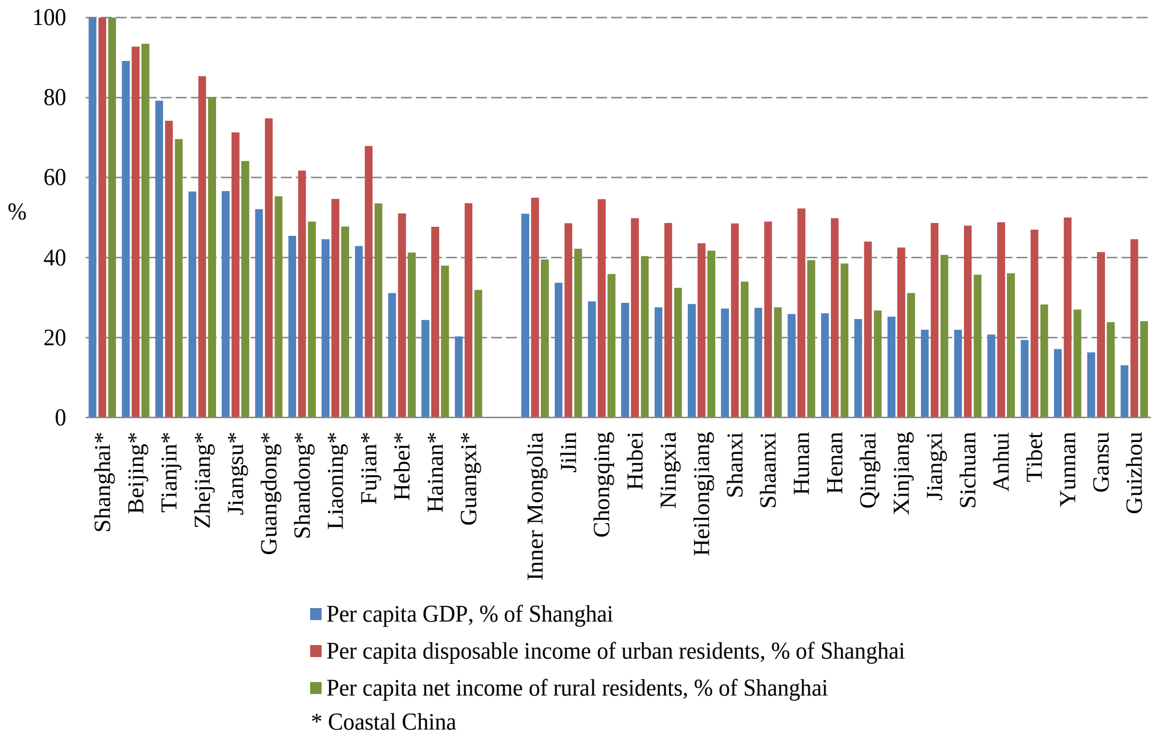

Figure 1.2. Comparison of the wealth of the population in China, 2009.

1.1.2. Health Care Reforms in China

China’s national health policy has also undergone several transitions along with the economic transition. The First National Health Conference, held in August 1950, laid down the following policies, which became hallmarks of the Maoist era: universal access to health care; prevention before cure; and integration of traditional and modern medicine (Mackerras, 2006). By the end of the 1950s, a three-tiered structure of health services provision came into being in both urban and rural areas. In addition, near-universal medical insurance coverage (Blomqvist, 2007; Yip and Hsiao, 2008) was implemented by the Cooperative Medical Scheme (CMS) in rural areas and the Government Insurance Scheme (GIS), Labor Insurance Scheme (LIS), and poverty aid program in urban areas (Ma *et al.*, 2008; Yip and Hsiao, 2008; Ramesh and Wu, 2009; Wang, 2011). Although the quality of health services was not particular high, the health care system at that time provided an affordable and equally accessible health care service for nearly all of the Chinese population (Wang, 2011).

However, just as with the other sectors of the national economy under economic planning, the health system also suffered from decreased efficiency in resource utilisation by state-owned hospitals (Bloom and Gu, 1997; Wang, 2009). Health care providers, including physicians, nurses, and administrators, were paid a fixed salary and in order to maintain public health care spending at an affordable level the prices of medication and services were set and adjusted periodically by the central government (Ma *et al*., 2008).

Even with these difficulties, the system produced a distinct improvement in basic health indicators, such as life expectancy and infant mortality (Blomqvist, 2007; Ramesh and Wu, 2009). These achievements were reached at a relatively low cost: total health expenditures formed only 3% of GDP in the early 1980s (Ramesh and Wu, 2009), an expenditure of U.S. $5 per capita for health care (Hsiao, 1995).

At the end of 1978, China started an unprecedented transition from a social-planning to a market-driven economic system with important consequences for the health care system and, consequently, health outcomes (Yip and Hsiao, 2008). Medical insurance schemes met severe challenges during this period. In rural China, the Cooperative Medical Scheme virtually ended following the collapse of the People's Commune System[4] in the late 1970s and early 1980s (Liu *et al.*, 1995; Dong, 2009). Compared to equality of access, the ideology of private responsibility was over-valued (Schiere, 2010). The principle set down by the government was that "the premiums were paid mainly by individuals themselves, supplemented by collectively pooled subsidies and supported by government policies" (Wang, 2011, page 283). The deficiencies in medical insurance funds management (e.g. corruption) further hindered the motivation to enroll in the government-run CMS program once it became voluntary in the early 1980s (Schiere, 2010). On the other hand, the rapid escalation of health care costs in urban China led in part to a fiscal crisis in both the GIS and the LIS; as a result workers were in effect uninsured (Yip and Hsiao, 1997). Alleviating the fiscal pressure became the theme of health care system reform (Wang, 2011). With reduced government subsidies to health care, and with more people in both rural and urban areas becoming uninsured, the key role of health care finance has been shifted to the individual (see Chapter 4 for details).

Government and enterprises struggled with the immense financial burden of these two systems (i.e. GIS and LIS), driving the pilot experiment of Medical Savings Accounts (MSAs) in 1994 (Yip and Hsiao, 1997; Dong, 2009), which later expanded to cover the whole country in 1998, named Urban Employee Basic Medical Insurance (UEBMI). This reform transformed urban Chinese health care from a comprehensive national welfare system to a salary-oriented social insurance plan (Dong, 2009). China was faced with the

[4] With the introduction of the Household Contract Responsibility System, the rural household, following the replacement of collectives, became the basic productive unit (Wang, 2011).

double-edged sword of having both a large uninsured population and rapid health care cost inflation (Yip and Hsiao, 2008). The large share of individual expenditure in national health spending has reduced accessibility and equality of health services. The lag between the rapidly growing economy and the rigid health system has posed severe challenges for the sustainability of health care financing and delivery.

Realising the problems, the country's policymakers began to implement a series of reforms aimed at correcting its perceived shortcomings in 2003 (Wagstaff *et al*., 2009b). Starting from rural areas, the New Rural Cooperative Medical Scheme (NRCMS) was proposed by the Ministry of Health of China, the Ministry of Finance of China and the Ministry of Agriculture of China nationwide to reestablish the formal CMS in 2003 (Ma *et al*., 2008; Dong, 2009). The new scheme highlights the responsibility of subsidy from all levels of government (Wang, 2011). In October of 2006, the Central Politburo orchestrated the creation of a 16-ministry National Health Care Reform Coordinating Committee, setting the stage for a nationwide debate on health system reforms. Like many other countries in the world, China has found health care reforms a challenging task. On July 10, 2007, the State Council issued a State Policy Document No. 20, on Urban Resident Basic Medical Insurance (URBMI), setting a landmark step towards the goal of establishing universal public health insurance coverage in China (Lin *et al*., 2009; Gu, 2009). This policy effort bears strong testament to the development goal of "building a people-oriented, harmonious society" that China has pursued since the Chinese Communist Party 16th Congress. By 2008 health care had become the leading public concern in China (Ho and Gostin, 2009).

China’s Health Care Reform Plan for 2009-2011, “Deepening the Health Care Reform” was unveiled in April 2009, which indicates the start of a new wave of health reform and represents a significant milestone in China’s path to establishing a strong national health system (Yip and Hsiao, 2009). Under the new plan, the government promised to contemplate four interrelated reforms of: the medical insurance system, public health, the medical service delivery system and the pharmaceutical supply system (Dong, 2009). A review by Yip *et al.* (2012) suggests that the reform is heading in the right direction; however, there are still four major deficiencies in the current health care system that need to be addressed: inequities, inadequate prevention and control of non-communicable diseases, and shortages of health professional resources. These issues, while important, are beyond the remit of this thesis and are left for future work.

1.1.3. Progress in Health in China

The health status of the Chinese population has improved along with the development of the economy and health care reform. Take life expectancy at birth (hereafter “life expectancy”), one of the most popular health indicators, as an example. Combining figures reported in Fang (1994), Riley (2004) and the numbers calculated from the National Population Census in the year 1982, 1990 and 2000 (National Bureau of Statistics of China, 2011), a clearly increasing trend of life expectancy can be found (see Table 1.2). There is also evidence suggesting a relatively larger improvement during the pre-reform period and a higher life expectancy for females. To compare the health achievement of China with other countries, data from a group of international organisations[5] can be used. According to the World Bank data, in 2009 the life expectancy of China (73.1 years) was well above the average level of middle income countries (68.8 years), slightly exceeding the East Asia and Pacific region level (73.0 years), but still 6.5 years behind high income countries (79.6 years).

Table 1.2. Life expectancy at birth of China, 1950-2000.

	1950	1957	1981	1990	2000
Life expectancy	40.1	57	67.8	68.6	71.4
Male	N/A	N/A	66.3	66.8	69.6
Female	N/A	N/A	69.3	70.5	73.3

Sources: 1950: Riley (2004); 1957: Fang (1994); 1981-2000: National Bureau of Statistics of China (2011).

Note: N/A: not available.

By using Stock charts and plotting the life expectancy at birth for each province in two time periods (1990 and 2000) (data are drawn from the National Population Census (National Bureau of Statistics of China, 2011)), a clear intra-province disparity can be observed, with a better health outcome in (relatively richer) coastal China[6] (see Figure 1.3 for details). The difference between the highest and lowest life expectancy at birth within mainland China was 15.3 years in 1990; the difference was almost the same in 2000, in that there was a 13.8 years difference between Shanghai and Tibet. Similar to other countries, females have a relatively higher life expectancy compared to males in all provinces, ranging from 1 year in Jiangxi (in 2000) to 6.4 years in Hainan (in 1990) (see

[5] Several international organisations have compiled the life expectancy data for the world, such as the World Bank, the WHO and the International Monetary Fund (IMF). These data should be used with caution as they may not be able to capture the real evolution of health status for a single country. Take China for example, comparing the predicted life expectancy from the World Bank and Banister (1992), there exists a huge difference in the early 1950s. See Appendix 1 for details.

[6] The first 12 provinces in Figures 1.3 to 1.5 refer to coastal China. The data for Chongqing is not present since it separated from Sichuan province in 1996.

Figures 1.4 and 1.5 for male and female, respectively). Not only do the life expectancies vary, other available indicators also suggest a similar intra-province disparity on health outcomes (see Chapter 5 for details).

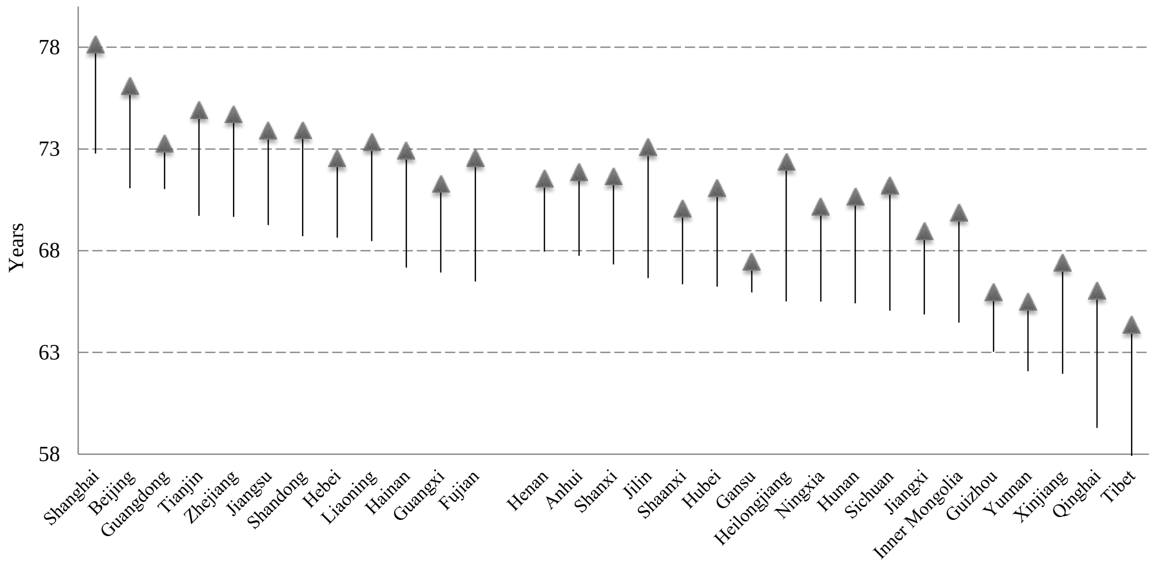

Figure 1.3. Life expectancy at birth of the Chinese population, 1990, 2000.

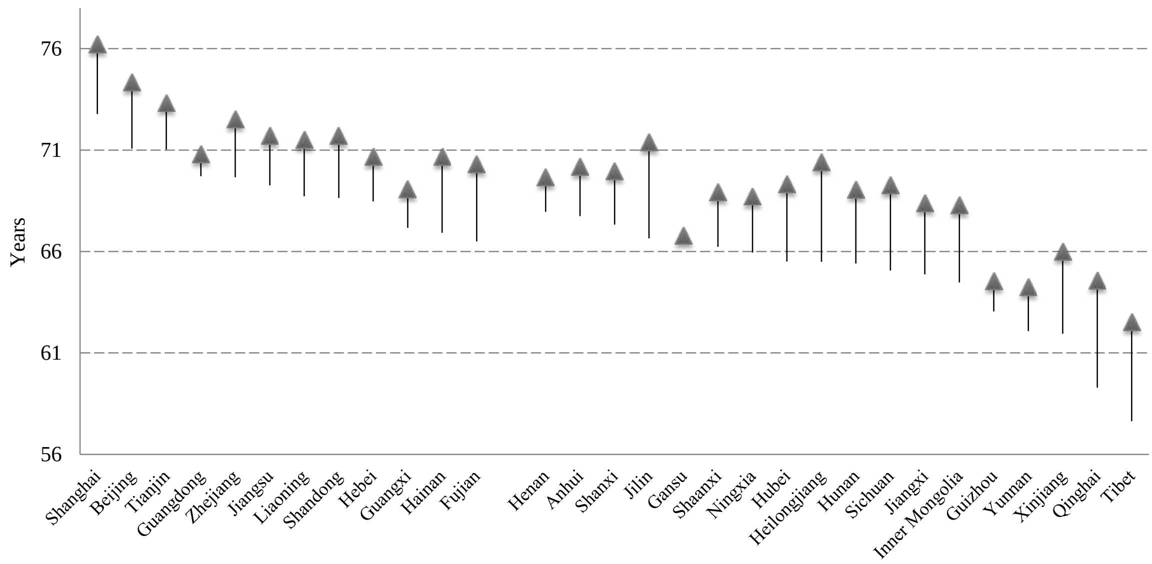

Figure 1.4. Male life expectancy at birth of the Chinese population, 1990, 2000.

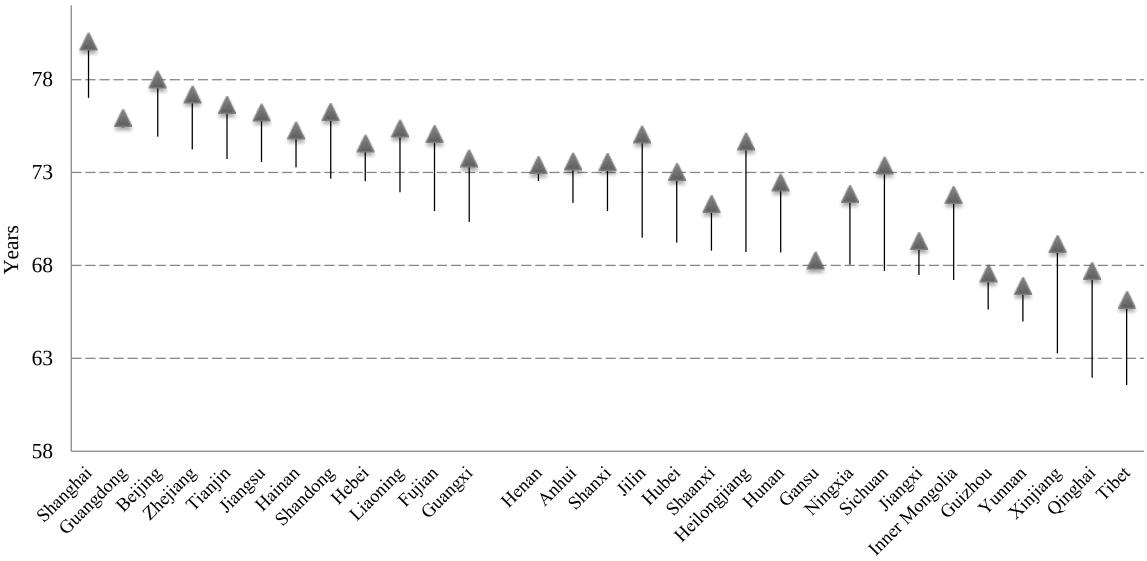

Figure 1.5. Female life expectancy at birth of the Chinese population, 1990, 2000.

The improvement of health status has also been documented by declining mortality rates. The crude death rate (CDR) is the only mortality indicator that is available continuously for over six decades. Corresponding to increased life expectancy at birth, the major decrease in CDRs happened in the pre-reform period. It can be clearly seen in Figure 1.6 that the decreasing trend was interrupted by the Great Leap Forward[7], and the following famine in the late 1950s and early 1960s (Kanbur and Zhang, 2005). Figure 1.6 also shows a disparity of rural-urban CDR, the CDR is higher in rural areas compared to urban China. More evidence and analysis about the rural-urban disparity on maternal and child mortality indicators can be found in Chapter 5.

[7] The Great Leap Forward refers to Chairman Mao Zedong's plan to develop agriculture and industry during the late 1950s and the early 1960s. For details of the health policy during the Great Leap Forward refer to Lampton (1974); for the demographic and economic consequences see Peng (1987) and Li (2000).

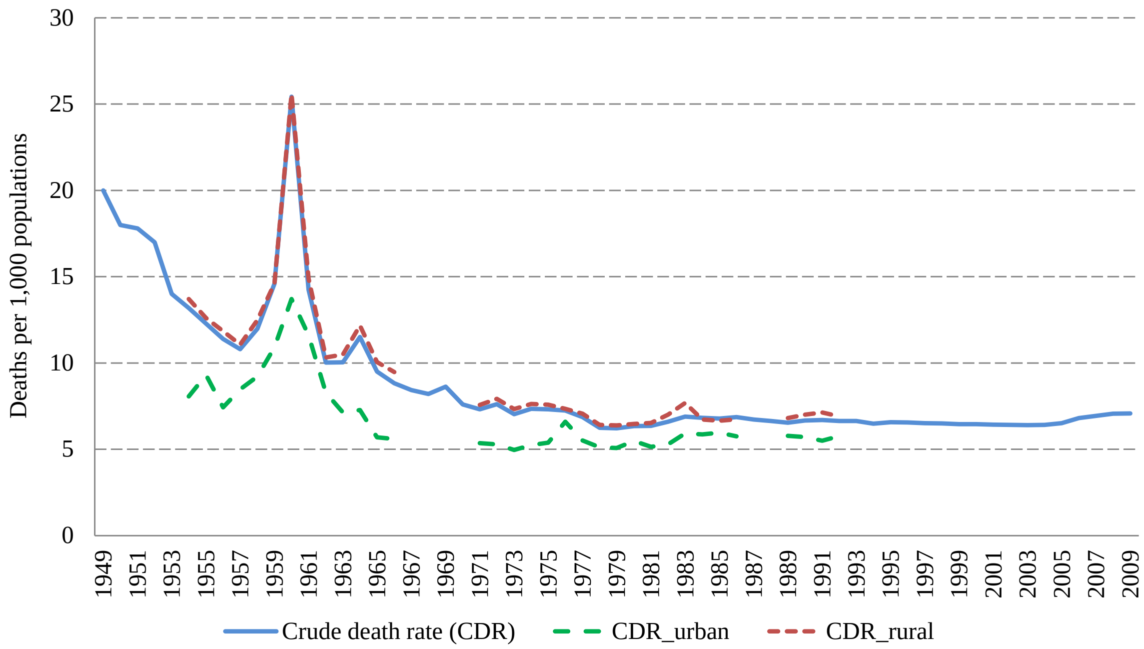

Figure 1.6. Crude death rate of the Chinese, 1949-2009.

The major causes of death reveal an important feature of the epidemiological transition in both urban and rural China. Figures 1.7 and 1.8 show the transition of seven major causes of death. The figures used here are based on the CDR, taken from Zhao (2007) and Ministry of Health of China (2010).

In urban China, infectious diseases (excluding tuberculosis) and pulmonary tuberculosis were among the top five major causes of death in 1957. The deaths owing to these two groups of diseases totalled around 15.4% of total deaths (Zhao, 2007), while the figure has dropped to 1.0% in the latest data. The top five major causes were malignant tumour, heart diseases, cerebrovascular diseases, respiratory diseases, and injury and toxicoses in 2009. See Figure 1.7 for details.

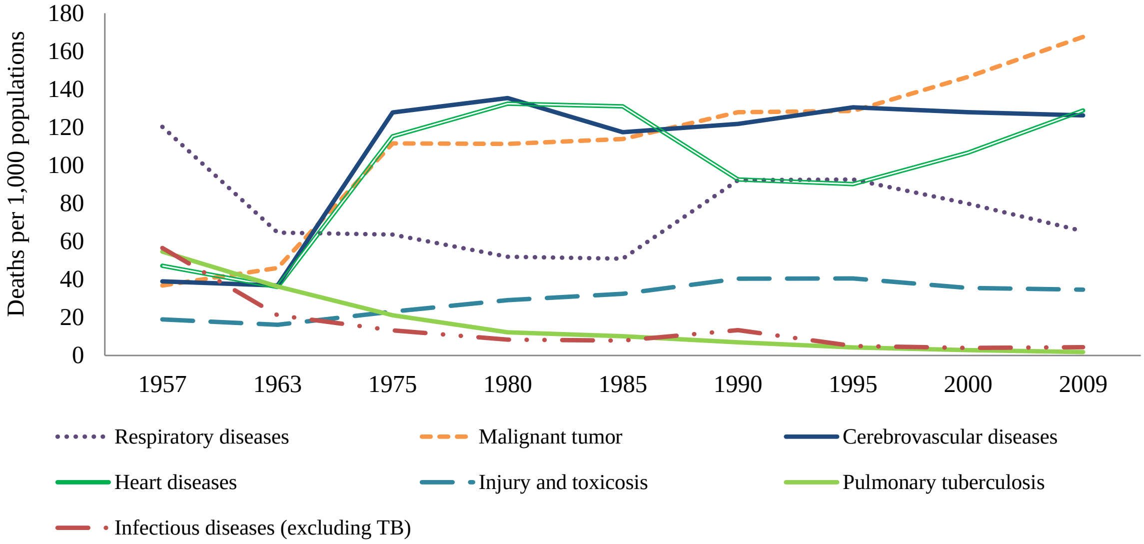

Figure 1.7. Disease-specific crude death rate in urban China, 1957-2009.

Similar to urban China, the top five major causes of death in rural China in 2009 are malignant tumour, cerebrovascular diseases, heart diseases, respiratory diseases, and injury and toxicoses. The data for infectious diseases (excluding tuberculosis) and pulmonary tuberculosis are missing before 1990; in 2009, they were ranked as 11 and 15 in the death causes, and responsible for only 1.1% of total deaths. See Figure 1.8 for details.

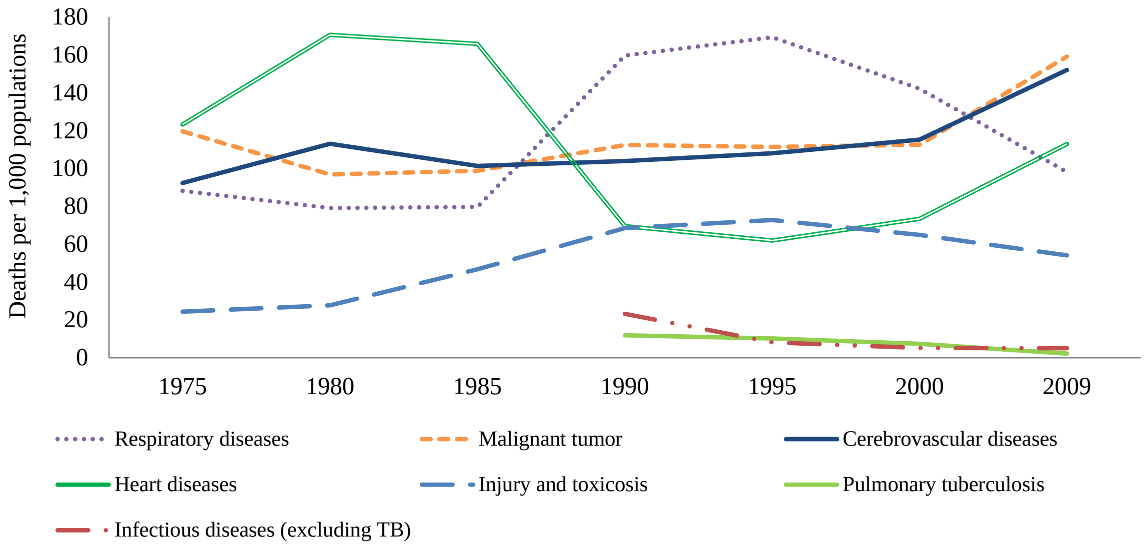

Figure 1.8. Disease-specific crude death rate in rural China, 1975-2009.

Self-reported health has also been found to provide a valid overall health outcomes assessment measure and is widely used in empirical studies (Idler and Benyami, 1997; Bopp *et al*., 2012). Using multiple-waves of the Medicare Current Beneficiary Survey, Lubitz *et al*. (2003) show that people aged 70 whose self-reported health status was excellent had a life expectancy of 13.8 years, 4.5 years longer than those who reported poor health status. Using summarised adults' self-reported health data from Zhao (2007), in which the self-reported health was calculated based on the China Health and Nutrition Survey (CHNS)[8] 2000, correlation analyses between self-reported health (the percentage of respondents who reported their health were good or excellent) and life expectancy at birth from nine provinces in the year 2000 are presented for urban and rural area separately in Figure 1.9, followed by two figures using gender sub-sample data.

If both health indicators reveal the same health status, it is expected that a positive correlation will be found between them. However, as shown in Figures 1.9 to 1.11, consistent negative correlations are reported. The negative correlation is stronger in urban China compared to rural China, and also stronger in the male sample compared to the female sample. The potential reason behind this finding is that in China, the two indicators capture different dimensions of health status. Organisation for Economic Co-operation and Development (OECD) also reports that the dispersion of country outcomes between adults' self-reported health status and life expectancy at birth is high, suggesting "these indicators capture separate dimensions of health status that do not necessarily move in tandem" (OECD, 2011, page 114).

[8] CHNS covers nine provinces in China. Within each province, a multistage, random cluster process was used to draw the samples. See http://www.cpc.unc.edu/projects/china for details. In CHNS 2000, self-reported health is coded in five levels: excellent, good, fair or poor, or very poor.

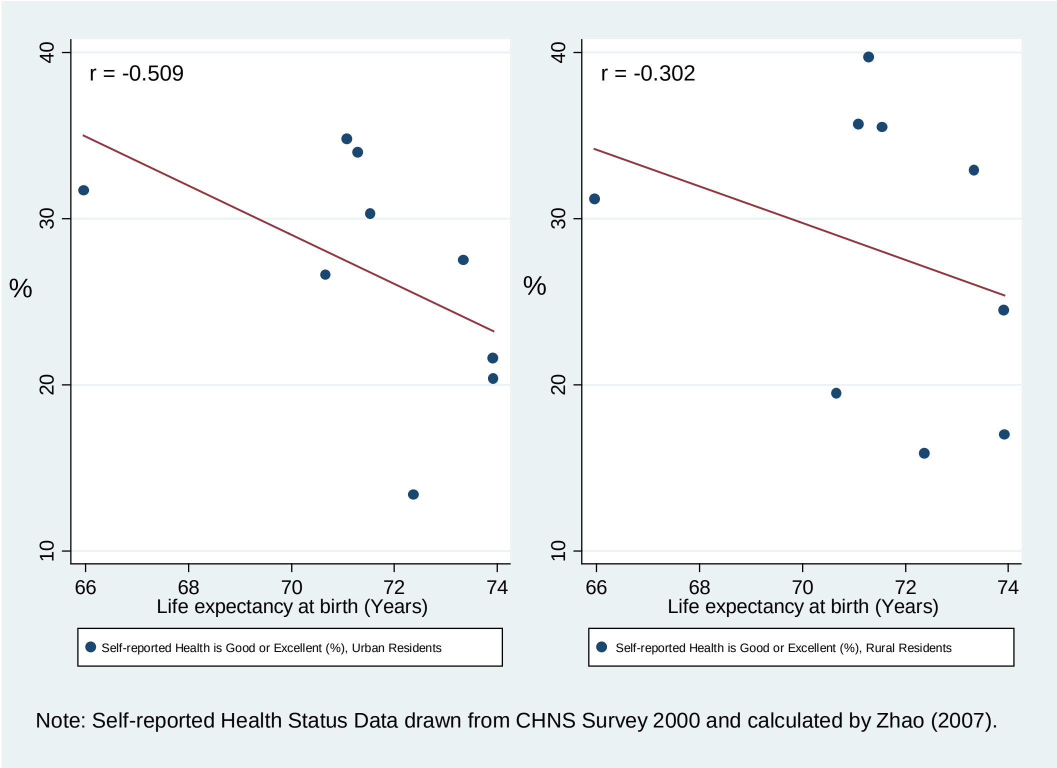

Figure 1.9. Correlation between self-reported health status and life expectancy at birth in 9 provinces, 2000.

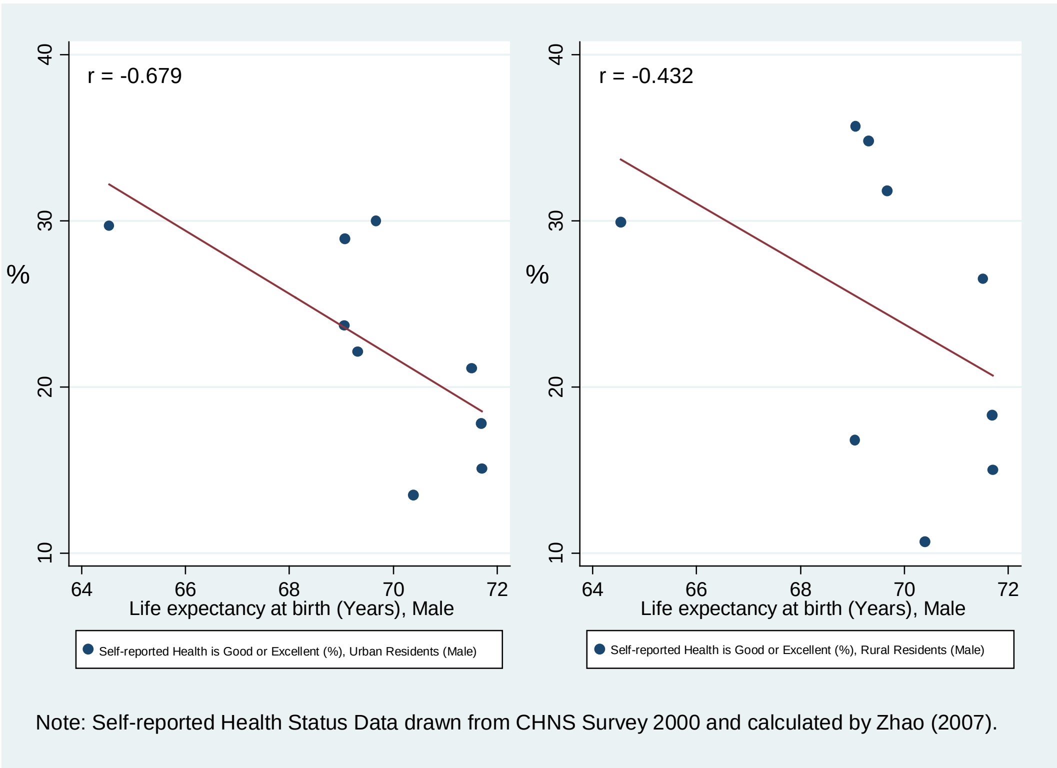

Figure 1.10. Correlation between self-reported health status and life expectancy at birth for male in 9 provinces, 2000.

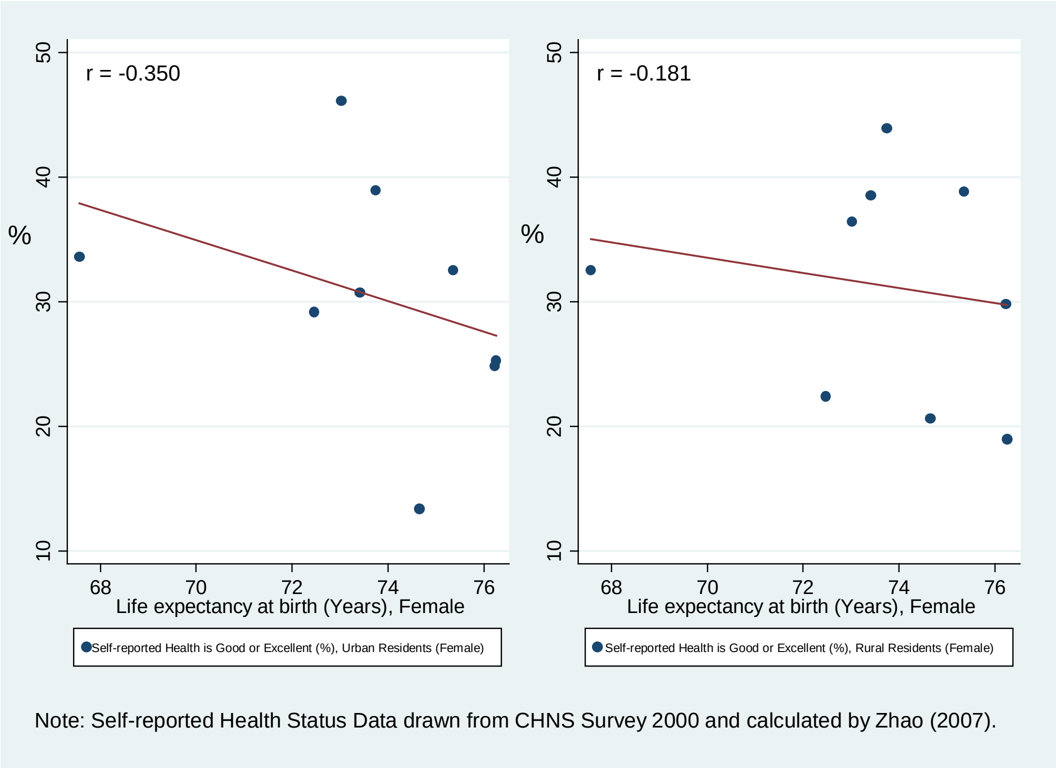

Figure 1.11. Correlation between self-reported health status and life expectancy at birth for female in 9 provinces, 2000.

In summary, along with economic transitions and health care system reforms, health status has witnessed a great improvement in prolonging life expectancy and a decline in the crude mortality rate. However, it should be noted that various indicators capture different dimensions of health status. China's development history provides a unique platform to study the crucial relationships among health investment, health outcomes and economic growth.

1.2. THESIS OBJECTIVES

The primary objectives of the thesis are as follows.

I. To investigate whether health investment has a significant impact on economic output in China using national time-series and provincial-level panel data.

II. To investigate the cyclical behaviour of public and private health expenditure in China using national time-series and provincial-level panel data.

III. To investigate the relationship between economic development, economic fluctuations and maternal and child health outcomes and infectious disease in China using national time-series and provincial-level panel data.

1.3. OVERVIEW OF THESIS

In Chapter 2, the relationship between health investment and GDP in China is investigated using health capital in a human capital-extended production function. The 45 years of (1967-2007) national level time series data are used. Considering the unknown nature of structural breaks (i.e. the presence of structural instability in a macroeconomic time series or a group of cointegration series), the Gregory-Hansen and Hatemi-J's cointegration tests were adopted to estimate breaks and the Fully Modified Least Squares, Canonical Cointegrating Regression and Dynamic Ordinary Least Squares were used to account for endogeneity and serial correlation. A structural break is found in 1976, when the Cultural

Revolution ends. A significant robust relationship between health investment and level of GDP in China over 45 years was found. Results suggest that sustained economic development will rely on further health investment.

Using a 29 year (1978-2006) panel of provincial level data from China, Chapter 3 further investigates the role of health capital in a human capital model of economic output. Robust evidence is found through panel cointegration analysis that health capital has a significant and positive effect on the Gross Regional Product in China; the effect being stronger in the inland regions compared to the coastal areas based on estimates that account for regional heterogeneity. Including potential structural breaks leaves the conclusions unchanged. This chapter highlights and discusses the potential role of diminishing returns to health investment in this globally important area.

Chapter 4 studies short-run cyclical behaviour of public (government and social) and private health expenditure and GDP using both time series and panel data techniques. First, national time series data (1978-2009) have been used within a multivariate Beveridge-Nelson decomposition framework to construct the permanent and cyclical components. The correlation analysis results for the cyclical components suggest that current public health expenditure is pro-cyclical (i.e. public health expenditure increases in good times and falls in bad times) while there is no clear evidence of a correlation between cycles in private health expenditure and in GDP growth. Next, using an instrumental variable method and the generalised method of moments estimator, provincial-level panel data (1997-2009) analyses confirm a pro-cyclical impacts of government spending on health. The provincial analysis also suggests that private health expenditure in urban China has a pro-cyclical association with GDP growth, but a lack of good instruments makes it difficult to identify a clear causal link between cycles in income growth and private health expenditure. The results suggest two policy recommendations relevant to public health expenditure, in line with China's current health reforms.

In Chapter 5, the relationship between economic development, economic fluctuations and several maternal and child health outcomes and also infectious disease epidemic are studied. Based on 19 years (1991-2009) of national time series data, correlation analyses are mainly used to study the relationship between four maternal and child mortality indicators and economic development in urban and rural China separately. It is found that an increased real income correlates with a low mortality rate. When using 8 years (2002-2009) of provincial-level panel data and controlling a group of relevant variables, it is found that increased real income and health services access or government health investment are significantly associated with an improved maternal and child health outcomes. However, there is evidence suggesting the improved real income is also significantly associated with an increased tuberculosis incidence in inland China. The potential association between economic fluctuation and health outcomes are also investigated. There is some evidence suggesting that economic fluctuation is significantly associated with perinatal mortality rate (based on the whole China sample) and tuberculosis incidence (in inland China).

Chapter 6 concludes the thesis by summarising key results and providing some policy suggestions to China and other developing countries in transition. The directions for future research are also discussed.

CHAPTER 2

HEALTH INVESTMENT AND ECONOMIC OUTPUT IN CHINA

2.1. INTRODUCTION

The traditional view of the nexus between health and income has emphasized the causal link from economic growth to improved health; however, with the recognition of the importance of health to productive human capital, the health-income relationship, has been explained by another causal link, running from health to income (Bloom and Canning, 2000). In other words, economic development is good for health, but good health also fosters economic growth (Suhrcke *et al*., 2007). Research on the role of health capital evolved from theory in the 1960s to empirical microeconomic studies in the 1970s, and to macroeconomic research in the late 1970s (e.g. Hicks, 1979). Alongside a substantial body of microeconomic evidence, Barro *et al*. and Bloom *et al*. began a burgeoning stream of empirical research in economic growth by looking into the role of health capital in economic growth from a macro perspective. A tabulated summary of empirical analyses investigating the effect of health on economic growth (output) can be found in Appendix 2.

The majority of empirical research has found a significant positive effect of health capital on economic growth, with the coefficients varying according to the countries analysed, the model specification, as well as the proxies used (see Appendix 2). One important finding from cross-country macro-empirical analyses is that health capital is a more significant factor for developing countries than developed countries (Knowles and Owen, 1995; McDonald and Roberts, 2002). However, insignificant effects have also been reported. Hartwig (2010) uses five-year interval panel data from 21 Organisation for Economic Co-operation and Development (OECD) countries (1970-2005) and rejects the hypothesis that health capital formation (proxied by life expectancy or health expenditure) is good for long-run economic growth in rich countries. Three explanations are offered: the growth effect of health capital formation is only short-lived (less than 5 years); the time period adopted in the study is too short to capture the health capital effect (using 13 OECD countries' very long-term data ranging from 1820-2001 to 1921-2001, Swift (2011) found that life expectancy has a significant positive effect on both Gross Domestic Product (GDP) and GDP per capita); or perhaps the effects are indeed zero. In related work, Webber (2002) uses calorific intake to proxy health capital and finds reducing under-nutrition will have a positive but insignificant effect on economic growth.

Few papers have explored the effect of health on economic growth in China, one of the largest developing countries. Li and Huang (2009) analysed the relationship between health capital stock (proxied by the number of doctors per 10,000 persons, or the number of hospital beds per 10,000 persons) and per capita GDP in an augmented Mankiw-Romer-Weil (MRW) model[9] based on Chinese provincial data from 1978-2005. Their results show that health capital has a positive and significant effect on China's GDP. Bloom *et al*. (2010) used a panel of countries over the period 1960 to 2000 (using five-year periods) to explain the take-off in growth in China and India; their results show that a rise in life expectancy has a significant positive effect on the growth rate of GDP per capita. Jamison *et al*. (2010) adopted a panel of countries over the period of 1960 to 2000 (with 10-year intervals) and by adding interaction terms between health and education capital with an indicator dummy variable for China, their empirical results suggest that the effect of health on growth in China is close to zero. On the other hand, Liu *et al*. (2008) from a micro perspective examined the economic returns at the household level from the health of its individual members using longitudinal data. Their study provided empirical evidence that household income is strongly influenced by the health of its members, particularly in rural areas.

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[9] Mankiw, Romer and Weil (1992) considered an augmented version of the Solow model, in which education was included as an input (similar to physical capital and raw labour) of the production function. Knowles and Owen (1995) further included health as an additional factor in the production function and it is referred to as an augmented MRW model here.

One important issue with cross-country panel data is heterogeneity, especially when a particular country may encounter significant structural breaks during the period examined. China provides a very good example, with its well documented economic history: the unprecedented transition from a social-planning to a market-driven economic system since the late 1970s. For example, based on analysis of China's annual GDP and its sectoral components data, Li (2000) found three significant breaks caused by the Great Leap Forward (1958-1960), the introduction of market reforms (1981-1983), and Deng's southern tour (1992). In addition the author points to two major episodes in the period 1952-1998: the Cultural Revolution (1966-1976) and macroeconomic austerity program (1990-1991). Smyth and Inder (2004) find, using provincial data for China, that for the majority of provinces the first break is associated with either the Great Leap Forward or the Cultural Revolution and the second break is in the market reform period.

This chapter further investigates the nexus between health investment and the level of GDP in China through a cointegration analysis, which allows for potential structural breaks. A set of econometric techniques that offer some distinct advantages over much of the previous literature have been adopted. First, in contrast to the previous literature on health and development, which is based on cross-country data, focusing on a single country like China can reduce problems with both cross-country heterogeneity and data consistency. Secondly, while much of the previous literature focuses on a bivariate relationship between health and GDP, this chapter estimates the long-run relationship under a production function framework where a standard set of inputs has been considered. The results will potentially shed light on the related theoretical models. Thirdly, when investigating a single country's economic performance, especially a country like China which has undergone significant transition, it is reasonable to assume that the cointegrating relationship may change at some point. Previous literature focuses on the issue of structural breaks in a single variable (e.g. GDP or health expenditure), and little emphasis has been put on the multivariate time series cointegration relationships with structural breaks.

The outcomes of this research have the potential to be quite important in the context of China's history of significant under-funding of health services. Examining empirically the effect of health investment on economic development will provide potentially useful evidence for those involved with policy development.

The remainder of this chapter is structured as follows: the empirical model, estimation methods and the data are presented in Section 2.2. Section 2.3 describes the results. Discussions and conclusions are summarised in Section 2.4.

2.2. METHODOLOGY AND DATA

2.2.1. Theoretical Model

The model is based on a production function extended to include human capital in two forms - education and health. The model assumes an aggregate production function with constant returns to scale (CRS).[10] The Cobb-Douglas production function with factor-augmenting technological process, for time period *t*, is:

$$Q_t = AF(K, E, H, L) = A_t^{\gamma} \cdot K_t^{\alpha} \cdot E_t^{\beta} \cdot H_t^{\eta} \cdot L_t^{\chi} \tag{2.1}$$

where $\chi = 1 - \alpha - \beta - \eta$. It is assumed that $\gamma, \alpha, \beta, \eta > 0$ and $0 < \chi < 1$. The notation adopted is as follows: *Q* is aggregate output, *K* is the stock of physical capital, *E* and *H* are variables that express the stock of education and health respectively, *A* is the level of technology, *L* is the level of aggregate employment or labour, and the subscript *t* denotes the time period. Assuming *A* grows exogenously at rate *g* over time:

$$A_t = A_0 e^{gt} \tag{2.2}$$

Taking the natural log of both sides of Equation (2.1) and combining with Equation (2.2) gives:

$$\ln Q_t = \alpha \ln K_t + \beta \ln E_t + \eta \ln H_t + \chi \ln L_t + \gamma \ln A_0 + \gamma g t \tag{2.3}$$

Define *q* as the level of output per worker, i.e., *q*=*Q*/*L*; similarly, define *k*, *e*, *h* as the stocks of physical capital stock, education capital stock and health capital stock per worker, respectively. Equation (2.3) can be written in 'per worker' form:

$$\ln q_t = \alpha \ln k_t + \beta \ln e_t + \eta \ln h_t + \gamma \ln A_0 + \gamma g t \tag{2.4}$$

For the sake of simplicity, rewrite Equation (2.4) into the following matrix form with an error term *u*:

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[10] The CRS assumption is tested by two steps: firstly, the Equation (2.3) is estimated using an ordinary least squares (OLS) estimator, and secondly, an *F*-test is adopted to see if the null hypothesis ($\alpha + \beta + \eta + \chi = 1$) holds. The result suggests the CRS assumption cannot be rejected (p = 0.117).

$$\ln q_t = \varsigma + \vartheta' x_t + \gamma g t + u_t \qquad (2.5)$$

where ς is an intercept ($\gamma \ln A_0$), $x_{it} = \left(\ln k'_{it}, \ln e'_{it}, \ln h'_{it}\right)'$. This equation provides the basic framework for the empirical model to be estimated.

2.2.2. Econometric Methodology

The econometric analysis aims to find the long-run effects of the independent variables through four steps. In the first step, unit root tests are used in order to see if the time series are stationary. In addition to the generalised least squares (GLS) augmented Dickey-Fuller (ADF) test proposed by Elliott *et al*. (1996), two other tests, which can deal with potential structural breaks, are adopted. This is important, as Perron (1989) points out that the power to reject the unit root null declines if one does not consider a possible known structural break in the trend function. Among several candidate tests, Perron and Vogelsang (1992) structural break unit root test, which allows for a single structural break, and the Clemente-Montañés-Reyes (Clemente *et al*., 1998) test that can perform unit root tests with double mean shifts are chosen. These two tests allow for two changes in the mean of the variable under the assumption of either an innovational outlier (IO, which allows for a gradual shift in the mean of the variable) or additional outlier (AO, which captures a sudden change in a series).

The second step of the analysis involves firstly using Pesaran *et al.* (2001)'s bounds testing procedure to test whether a group of variables are cointegrated. This test can be applied irrespective of whether the regressors are $I(0)$ or $I(1)$. Microfit Version 5.0 software is used to perform the test. In this study, if the calculated F-statistics or W-statistics exceed the upper bound of the critical value band for $I(1)$ series, the null of no cointegration can be rejected. As mentioned in the Introduction, several potential structural breaks may exist, given China's economic development. To allow for a more general type of cointegration without the prior knowledge of the potential break dates, the test developed by Gregory and Hansen (hereafter GH) (1996) and extended by Hatemi-J (2008) is considered. The GH test is performed by "gregoryhansen.src" procedure using RATS Version 7.1 software; the Hatemi-J's test is performed by "CItest2b.prg" procedure (Hatemi-J, 2009) using GAUSS Version 9.0 software. The GH approach allows for a regime shift in either the intercept or the entire coefficient vector while the latter approach allows for two regime shifts on both the intercept and the slopes. That is, the GH test allows for cointegration with structural change in three forms: the level shift model (C model), both level shift with the trend model (C/T model), and the regime shift model (C/S model, both a level and slope change). The most relevant test to this study is the one based on the C/T model since in the main equation, Equation (2.4), there is a time trend[11] to capture the effect of technology growth. The test results based on C/S model are also reported. The C/T model can be expressed in Equation (2.6):

$$\ln q_t = \varsigma_0 + \zeta\varphi_t + \vartheta' x_t + \gamma g t + u_t \tag{2.6}$$

where ς_0 represents the intercept before the shift and ζ represents the change in the intercept at the time of the shift. φ_t is the structural break dummy defined as:

[11] From an econometrics perspective, Wooldridge (2009) also suggest it would still be a better choice to include a time trend into the model, to account for the trending behavior of the variables.

$$\varphi_t = \begin{cases} 0 \; if \; t \leq [n\tau], \\ 1 \; if \; t > [n\tau], \end{cases} \tag{2.7}$$

where the unknown parameter $\tau \in (0,1)$ denotes the relative timing of the break point, and [ ] denotes integer part. The C/S model can be expressed in Equation (2.8):

$$\ln q_t = \varsigma_0 + \zeta\varphi_t + \vartheta_0' x_t + \vartheta_1' x_t \varphi_t + u_t \tag{2.8}$$

where ϑ_0 represents the cointegrating slope coefficients before the regime shift, ϑ_1 represents the change in the slope coefficients. For each possible regime break point, the ADF test statistic is computed. The smallest value of the test statistic across all possible break points is chosen. Based on a similar idea, Hatemi-J (2008) extends the C/S model in the following way:

$$\ln q_t = \varsigma_1 + \zeta_1\varphi_{1t} + \zeta_2\varphi_{2t} + \vartheta_0' x_t + \vartheta_1' x_t \varphi_{1t} + \vartheta_2' x_t \varphi_{2t} + u_t \tag{2.9}$$

where φ_{1t} and φ_{2t} are structural break dummies defined as:

$$\varphi_{1t} = \begin{cases} 0 \; if \; t \leq [n\tau_1], \\ 1 \; if \; t > [n\tau_1], \end{cases} \tag{2.10}$$

and

$$\varphi_{2t} = \begin{cases} 0 \; if \; t \leq [n\tau_2], \\ 1 \; if \; t > [n\tau_2], \end{cases} \tag{2.11}$$

with the unknown parameters $\tau_1 \in (0,1)$ and $\tau_2 \in (0,1)$ denoting the relative timing of the break point.

The third dimension to the analysis is to allow for the possibility of simultaneity between the capital variables and economic output. This is done by estimating cointegrating long-run relationships for empirical linear models using three approaches: fully modified least squares (FMLS) (Phillips and Hansen, 1990), canonical cointegrating regressions

(CCR) (Park, 1992), and Dynamic Ordinary Least Squares (DOLS) (Stock and Watson, 1993). The FMLS approach applies a nonparametric correction to take account of the impact on the residual term of autocorrelation and possible endogeneity. The CCR procedure is also a nonparametric method quite similar to FMLS; however, in contrast to FMLS, which transforms both data and estimates, the CCR method only involves data transformations. The DOLS estimator is a semi-parametric approach involving adding leads and lags of differenced explanatory variables to a static cointegration regression to eliminate the potential bias due to endogeneity or serial correlation. In this case, the statistical inferences for the coefficients based on heteroskedasticity and autocorrelation consistent (HAC) standard errors are valid. For optimal estimation of various covariance matrices, the approaches of Andrews and Monahan (1992), and Christou and Pittis (2002) are followed, with the adoption of prewhitened quadratic spectral (QS) kernel estimators, in which the bandwidth parameter is selected via the nonparametric procedure of Newey and West (1994). This method tends to improve the accuracy of test statistics. Monte Carlo evidence regarding the relative finite sample behavior of the three estimators can be found in Inder (1993), Montalvo (1995), Cappuccio and Lubian (2001), Christou and Pittis (2002) and Kurozumi and Hayakawa (2009); the results are, however, inconclusive. Hence cointegration results from all three estimation methods are reported to assess the robustness of the findings. All cointegration analyses are performed by EViews Version 7.0 software.

The direction of causality between health capital and economic output potentially runs in both directions, as discussed in Chapter 1 and the Introduction. In this chapter, under the production function framework and the adoption of cointegration analysis, health capital is specified as an explanatory variable and it is tested to see whether it has a significantly positive effect on economic output. Since the direction of causality cannot be inferred from a cointegration equation, in the final step, a multivariate vector error-correction (VEC) model is estimated to test for Granger causality (Engle and Granger, 1987). Specifically, two types of causality tests, the short-run Granger non-causality test (whether the lagged dynamic terms are significant or not), and the weak exogeneity test (whether the one-period lagged error-correction term, ECT_{t-1}, derived from the long-run cointegration equation is significant or not; also called long-run Granger non-causality test) are considered (Singh, 2008).

2.2.3. Data and Variables

National time series data are used in this chapter. Publicly available Chinese annual data containing health investment information can be traced back to 1949 and other key variables can be found as early as 1952. However, there are concerns about the quality of the economic output data (i.e. GDP) prior to 1963 (see Wu, 2002), so the empirical analysis starts from 1963. Dernberger (1980) provides a detailed discussion of problems with China's output data. In a similar vein, Wu (2002) argues that the data in the Great Leap Forward years (1957-60) are most controversial and sharp fluctuations can be found during 1958-62. Further, considering 1962 was the end of China's Second Five-Year Plan period (1958-1962), and given Premier Zhou's proposal for national economic adjustment in February 1962, it seems prudent to use data from 1963 onwards. Considering the limited time period, it is appropriate to avoid another potential big exogenous shock (i.e. the Global Financial Crisis starts in 2008) at the end of the data. Hence, the chosen data range from 1963 to 2007 (45 years).

The output variable is measured by real GDP (constant 1978 Yuan[12]), defined as nominal GDP divided by the GDP implicit price deflator[13] in each year. Real physical capital stock (in constant 1978 Yuan) is estimated by the standard perpetual inventory method (PIM) following Wang and Yao (2003), while the initial aggregate capital stock in 1952 is set at 2,213 billion Yuan (in constant 1978 Yuan) (calculated by Chow, 1993 and used by Chow and Lin, 2002; Wang and Yao, 2003). In these calculations, real fixed capital investment is adjusted by an investment deflator, and the overall depreciation rate is 4% (adopted by the World Bank, 1997; Chow and Lin, 2002). The calculated real physical capital stock in this chapter is highly correlated with those reported in Wang and Yao (2003) (r = 1.000) and Chow and Lin (2002) (r = 1.000). The number of employed persons at year-end is used as the level of employment. Unless specified, all data are drawn from the National Bureau of Statistics of China (2005, 2008).

The stock of people with education at regular institutions of higher education level (*EDU*)

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[12] The difference between two popular names of China's currency (Yuan vs. RMB) is described here: http://www.bbc.co.uk/news/10413076.

[13] The GDP implicit price deflator is available from United Nations (UN) DATA (2009). An alternative way to calculate real GDP is to combine the nominal value of GDP and the GDP index. The real GDP calculated by those two approaches are identical. See Zhang (2006) for details.

is used to proxy education capital stock. As noted by the OECD, "flows of university graduates are an indicator of a country's potential for diffusing advanced knowledge and supplying the labor market with highly skilled workers" (OECD, 2003, page 50). It is calculated following Wang and Yao (2003)'s Equation (3.5) as:

$$EDU_t = (1 - \Omega_t)EDU_{t-1} + CG_t \tag{2.12}$$

where Ω_t and CG_t are the mortality rate of the population and the number of graduates at regular institutions of higher education in year *t*, respectively. At the end of 1951, 0.4%[14] of the whole population (563 million people) had higher education qualifications – this is used to calculate the initial value for *EDU*. It is not ideal to assume the mortality rate is independent of the level of schooling attained; however, this assumption is more realistic when data are scaled. Also, although there may be a gap between this proxy and the real education capital stock, this proxy captures the movement of education capital stock accumulation well. Using a shorter provincial-level panel data (1997-2006) analysis, Zhang and Zhuang (2011) found that tertiary education plays a more important role than primary and secondary education in explaining economic growth in China. A review and an empirical study highlighting the effect of higher education human capital on economic growth in developing countries has been undertaken by Gyimah-Brempong *et al*. (2006).

The effect of this variable is compared with another, the average years of schooling per worker; when the sample is restricted to start in 1978 there is sufficient data to calculate

[14] According to Barro and Lee (1996), for the population aged 15 and over in all developing countries, 0.4% had full higher education in 1960; the percentage was 0.8% for East Asia and the Pacific, 0.1% for South Asia.

the average years of schooling per worker. Results show that those two variables have almost the same effect in this case (see Table 2.11 for details). Following Qian and Smyth (2006), the average years of schooling (AYS_t) per worker is calculated as follows:

$$AYS_t = 1E_{1t} + 5E_{2t} + 8E_{3t} + 11.5E_{4t} + 14.5E_{5t} \tag{2.13}$$

where E_1 to E_5 stands for the different proportions of illiterate and semiliterate, elementary school, junior high school, senior high school and, college and above in time periods t. The lengths of schooling cycles for the above categories are assumed to be 1, 5, 8, 11.5 and 14.5 respectively. The above lengths are set following Wang and Yao (2003), except for the illiterate and semiliterate group for which this study assumes some of these people may have some years of formal education. The calculated average years of schooling per worker in this chapter is highly correlated with the one reported in Wang and Yao (2003), with a simple correlation of 0.991. The data of the educational attainment composition of the employed are drawn from Tan (2007) and the National Bureau of Statistics of China and Ministry of Labour and Social Security (1999-2008).

The proxy variables for health capital in macroeconomic empirical research are generally classified into two groups: health outcomes and health inputs. Health outcomes are characteristics determined both by an individual's health inputs and by their genetic endowment (Weil, 2007). For example, the most widely used is life expectancy, see Bloom *et al*. (2004). Health inputs are the physical factors that influence an individual's health (Weil, 2007) such as: health expenditure (Hartwig, 2010), health resources (Li and Huang, 2009), and nutritional intake (Webber, 2002).

For this study, since high quality health outcome data is unavailable in China, the variable chosen is similar to Li and Huang (2009), and Benos and Karagiannis (2010) - the number of doctors per 10,000 employed persons, to proxy for per worker health capital stock. Webber (2002) also suggests the number of doctors and nurses per capita, access to clean water could be adopted to investigate the effect of health-related areas on economic growth. Health care has been regarded as an important prerequisite for accelerating economic development (Fang, 1994), and the doctor is the principal health care professional. Or *et al*. (2005) highlights that physician numbers are a key determinant of mortality across OECD countries; similarly, Banister and Zhang (2005) using province level data from China also found that increased number of doctors is a significant determinant for several maternal and child health outcomes. Suhrcke and Urban (2010) includes the number of doctors per population in a growth model as a proxy for the health system inputs into the production of health. Li and Huang (2009) note the proxy variable used here focuses on the size of the highly skilled health workforce, and hence provides some indication of both the quantity and quality of health investment. One additional advantage for this proxy is that the variable itself actually captures the accumulation of past health investment considering the training time for health professionals (e.g. undergraduate clinical medicine training in mainland China takes five years and it may take at least one more year for an intern to be formally licensed). In this sense, this health capital proxy can be treated as predetermined and hence less sensitive to endogeneity.[15] Furthermore, adopting this proxy can avoid a

[15] Health expenditure or the number of hospital beds will be more likely to have a simultaneous effect with economic output.

potential bias of using life expectancy - that if life expectancy is high, there will be an older work force with higher levels of experience.

Summary statistics for the dependent and explanatory variables are presented in Table 2.1.

Table 2.1. Summary statistics.

Variable	Mean	Std. Dev.	Min	Max	Obs.
ln(GDP per worker)	7.310	0.794	6.125	8.869	45
ln(physical capital per worker)	8.242	0.809	7.244	9.954	45
ln(education capital per worker)	4.962	0.374	4.621	6.034	45
ln(health capital per worker)	3.257	0.113	2.995	3.436	45
ln(average years of schooling per worker)	1.841	0.143	1.552	2.039	30

2.3. RESULTS

2.3.1. Unit Root Test Results

The detailed results of unit root tests (with or without considering potential structural breaks) are reported in Tables 2.2-2.4. Different tests sometimes suggest a different order of integration for the series. There are alternative positions on which group of test results is the most accurate. From an economic perspective, some policies have a gradual effect on a series which suggest the IO tests should be focused upon; however, from a statistical perspective, sometimes the movement of data (one example, is a big jump of the number of employed persons at year-end from the year 1989 to 1990, see Cai and Wang (2010) for details) is also because of the changed statistical criteria by the National Bureau of Statistics of China which means the AO tests would be suitable.

The priority selection rule, as adopted by Madsen *et al*. (2012), suggests that comparing the three circumstances (no-break, one break and two breaks), the one break case should be preferred to the no-break case if the break is statistically significant and the two break case should be preferred if the second break is statistically significant. According to this rule, all series are proved to be non-stationary and integrated of order one.

Table 2.2. The GLS augmented Dickey-Fuller unit root tests results, 1963-2007.

	Lags	Statistics	Trend
Level form			
ln(GDP per worker)	2	-0.506	Yes
ln(physical capital per worker)	0	1.037	Yes
ln(education capital per worker)	0	1.405	Yes
ln(health capital per worker)	0	-1.531	Yes
ln(average years of schooling per worker)	0	-0.583	Yes
First difference form			
Δ ln(GDP per worker)	0	-3.828**	No
Δ ln(physical capital per worker)	0	-2.126	No
Δ ln(education capital per worker)	1	-0.996	No
Δ ln(health capital per worker)	0	-5.210**	No
Δ ln(average years of schooling per worker)	0	-2.989**	No
Second difference form			
Δ^2 ln(physical capital per worker)	0	-8.186**	No
Δ^2 ln(education capital per worker)	0	-9.676**	No

Notes: ** indicates the significance level at 5%. The number of lags (with a maximum of 5) to be included was selected by Schwarz Bayesian Criterion. Whether a linear trend is included (Yes: included; No: not included) for each variable is decided by economic theory and line graph. The 95% critical values for unit root tests are computed by simulations using 20,000 replications.

Table 2.3. Perron-Vogelsang unit root tests results (one break, 1963-2007).

Variable	Lags	TB_1	DU_1	(ρ-1)
		Innovational Outlier Case		
Level form				
ln(GDP per worker)	2	1990	0.037	0.002
ln(physical capital per worker)	1	1990	0.036**	-0.018
ln(education capital per worker)	1	1990	0.013	0.043
ln(health capital per worker)	0	1974	0.066**	-0.244
ln(average years of schooling per worker)	5	1997	0.017**	-0.149
First difference form				
Δ ln(GDP per worker)	1	1989	0.040**	-1.075**
Δ ln(physical capital per worker)	0	1989	0.037**	-0.619**
Δ ln(education capital per worker)	1	1989	0.027**	-0.302
Δ ln(health capital per worker)	0	1970	0.049**	-0.943**
Δ ln(average years of schooling per worker)	0	2003	-0.021**	-0.958**
		Additional Outlier Case		
Level form				
ln(GDP per worker)	0	1996	1.469**	-0.141
ln(physical capital per worker)	0	1999	1.599**	-0.119
ln(education capital per worker)	0	1999	0.819**	-0.181
ln(health capital per worker)	0	1977	0.184**	-0.273
ln(average years of schooling per worker)	5	2000	0.228**	-0.064
First difference form				
Δ ln(GDP per worker)	1	1988	0.020	-1.083**
Δ ln(physical capital per worker)	2	1988	0.043**	-0.233
Δ ln(education capital per worker)	2	1988	0.053**	-0.126
Δ ln(health capital per worker)	0	1969	0.048**	-0.939**
Δ ln(average years of schooling per worker)	2	2002	-0.014**	-0.698

Notes: Δ is the first difference operator. TB_1 is the time period when the mean is being modified; DU_1 is a sustained dummy variable capturing a mean shift occurring at the break date; ρ is autoregressive parameter. ** indicates the significance level at 5%. Lags included were selected following the procedure suggested in Clemente *et al*. (1998) with a max lag of 5.

Table 2.4. Clemente-Montañés-Reyes unit root tests results (two breaks, 1963-2007).

Variable	Lags	TB_1	TB_2	DU_1	DU_2	(ρ-1)
Innovational Outlier Case						
Level form						
ln(GDP per worker)	0	1967	1990	0.024	0.085**	-0.038
ln(physical capital per worker)	1	1990	2000	0.041**	0.032	-0.037
ln(education capital per worker)	1	1990	2001	0.036**	0.058**	-0.030
ln(health capital per worker)	0	1975	2000	0.084**	-0.045**	-0.315
First difference form						
Δ ln(GDP per worker)	1	1975	1989	0.028	0.024	-1.133**
Δ ln(physical capital per worker)	0	1968	1989	0.035**	0.036**	-0.745**
Δ ln(education capital per worker)	0	1980	1989	0.026**	0.021**	-0.549**
Δ ln(health capital per worker)	0	1970	1982	0.086**	-0.042**	-1.114**
Additional Outlier Case						
Level form						
ln(GDP per worker)	0	1981	1996	0.904**	0.964**	-0.352
ln(physical capital per worker)	0	1981	1997	0.872**	1.090**	-0.313
ln(education capital per worker)	1	1990	2000	0.403**	0.550**	-0.442
ln(health capital per worker)	0	1977	2003	0.194**	-0.079**	-0.335
First difference form						
Δ ln(GDP per worker)	3	1974	1988	0.015	0.013	-1.297**
Δ ln(physical capital per worker)	1	1969	1988	0.048**	0.032**	-0.740**
Δ ln(education capital per worker)	1	1988	2003	0.037**	0.079**	-0.747
Δ ln(health capital per worker)	0	1971	1981	0.073**	-0.039**	-1.110**

Notes: Δ is the first difference operator. TB_1 and TB_2 are the time periods when the mean is being modified; DU_1 and DU_2 are sustained dummy variables capturing a mean shift occurring at the break date; ρ is autoregressive parameter. ** indicates the significance level at 5%. Lags included were selected following the procedure suggested in Clemente *et al*. (1998) with a max lag of 5.

2.3.2. Cointegration Results

Bounds tests are firstly used to see if there is a cointegrating relationship. As can be found in Table 2.5 (Panel C), both *F*-statistics and *W*-statistics exceed the upper bound of a 5% critical value band, thus the null hypothesis of no cointegration is rejected. To explore the potential for a more general type of cointegration, Gregory-Hansen cointegration tests are adopted to see if there is a co-integrating relationship, allowing for one structural break at an unknown point. Results from Table 2.5 (Panel A) suggest that the null hypothesis of no cointegration for the C/S model (regime shift model) cannot be rejected; however there is evidence suggesting cointegration with the C/T model (level shift with the trend model) at the 5% level. The evidence in Figure 2.1 suggests a break at 1976.[16]

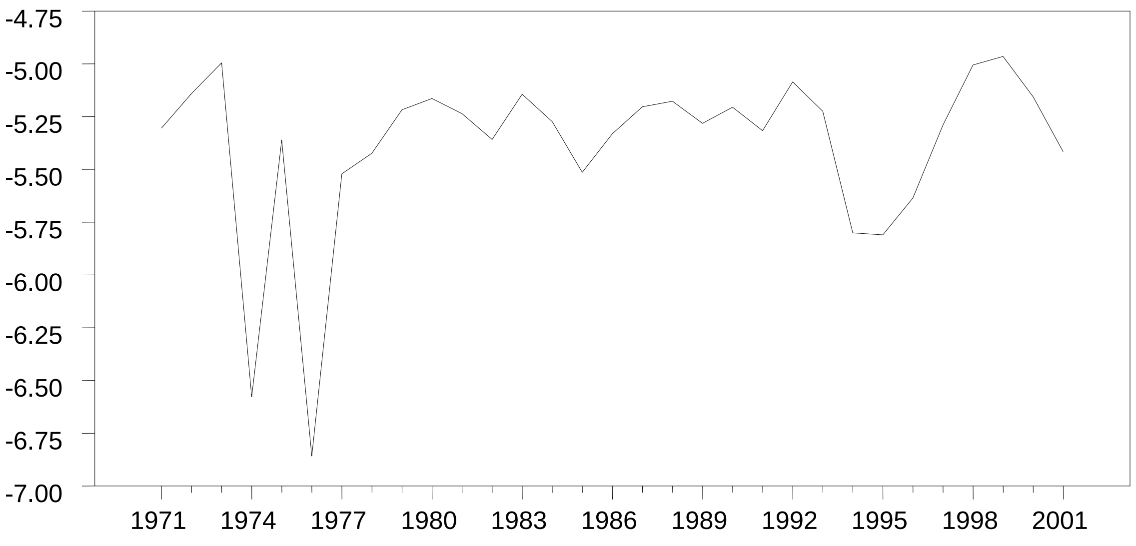

Figure 2.1. Gregory-Hansen cointegration tests, 1963-2007.

[16] As noted in Narayan (2005), the difference between the identified break dates by unit root tests and cointegration test is mainly because the former one searches for breaks in a series while the later one searches for breaks in the residual of testing series.

Table 2.5. Cointegration tests results, 1963-2007.

Method	Test	Statistics	Critical Values	
			5%	10%
Panel A: Residual based test for cointegration by Gregory and Hansen (1996)				
C/T model	ADF*	-6.858	-5.57	-5.33
C/S model	ADF*	-5.304	-6.00	-5.75
Panel B: Residual based test for cointegration by Hatemi-J (2008)				
C/T model	ADF*	-5.667	-7.35	-7.12
	Z_t*	-5.092	-7.35	-7.12
	Z_α*	-35.532	-104.86	-97.75
Panel C: Bounds test for cointegration by Pesaran et al. (2001) (intercept and linear time trend case)				
No break	*F*-statistics	12.716	5.57	4.76
	W-statistics	50.864	22.26	19.02
One break identified by Gregory and Hansen (1996)	*F*-statistics	10.195	6.10	5.23
	W-statistics	40.778	24.40	20.90

Notes: C/T model: Level shift with trend; C/S model: Regime shift. For Gregory and Hansen (1996) and Hatemi-J (2008) cointegration tests, lags included were selected by truncation at longest lag with P-value < 0.1; critical values are drawn from cited papers. For bounds tests, the lag is selected based on Schwarz Bayesian Criterion. Only critical values of upper bound for *I*(1) regressors have been reported; the critical value bounds are computed by stochastic simulations using 20,000 replications.

Hatemi-J's two regime shift test results, reported in Panel B, further suggest that the null hypothesis of no cointegration for the two regime shift model also cannot be rejected. Also included is the one structural break dummy identified by Gregory-Hansen approach (C/T model) and the bounds test is undertaken again; the results again confirm that the null of no cointegration can be rejected. The importance of whether to include the identified structural break dummy could also be guided by the history of the development of the economy, this will be discussed in detail later. Basically, it is concluded that there is one significant intercept break that should be accommodated in the model.

Before presenting the key cointegration results (in Table 2.7), the simple correlation analysis of variables used in the cointegration equation is shown in Table 2.6. There is a very high correlation between ln(GDP per worker) and ln(physical capital per worker). This is also documented in Wang and Yao (2003) who found a very similar calculated correlation of 0.994 (for the years 1963-1999).

Table 2.6. Simple correlation statistics, 1963-2007.

	ln(GDP per worker)	ln(physical capital per worker)	ln(education capital per worker)	ln(health capital per worker)
ln(GDP per worker)	1.000			
ln(physical capital per worker)	0.996	1.000		
ln(education capital per worker)	0.918	0.937	1.000	
ln(health capital per worker)	0.469	0.460	0.263	1.000

Estimates of the cointegration regressions using the 1963-2007 data are shown in Table 2.7. As shown, without allowing for the potential structural break, aside from the DOLS estimator, the other three estimators show that health capital has a positive but insignificant effect on GDP; when including the dummy variable, which equals 1 during the years after the break - 1976 (Columns (2), (4), (6) and (8) in Table 2.7), a new picture emerges. The three efficient estimators (i.e. FMLS, CCR and DOLS) all find health capital has a significant and positive effect on GDP at the 10%, or 5%, significance level. With the DOLS estimator, the number of leads and lags included are changed to see if the results are sensitive to this. Table 2.8 shows that the results are robust for health capital - the significant positive coefficients range from 0.397 to 0.568. The different coefficients on health capital from the three estimators reported in Tables 2.7 and 2.8 suggest a 10% increase in health significantly contributes to around 2.52% to 5.68% increase in GDP, indicate a robust and strong positive effect.

To check if the above significant positive effects of health capital on GDP above are robust, employed persons are replaced with the size of the population. In this way, the per capita form of Equation (2.4) is estimated. Again, the Gregory-Hansen cointegration test shows that there is a significant structural break in 1976, at 1% significance (the C/T model, *t*-statistic = -6.685); the estimates of the long-run cointegrating vectors suggest that health capital has a significant and positive effect on GDP (ranging from around 0.302 to 0.460, depending on the choice of three efficient estimators; see Table 2.9 for details).

Results from Tables 2.8 to 2.10 suggest education capital has a significant positive effect on economic output in China; however, although physical capital is positive, it is insignificant.

Table 2.7. Estimations of the long-run cointegrating relationship, 1963-2007.

	OLS		FMLS		CCR		DOLS	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(physical capital per worker)	0.310*	0.325**	0.225*	0.105	0.229***	0.134	0.067	0.167
	[0.164]	[0.159]	[0.113]	[0.172]	[0.080]	[0.110]	[0.197]	[0.219]
ln(education capital per worker)	0.312**	0.243*	0.361***	0.445***	0.375***	0.476***	0.720***	0.586**
	[0.145]	[0.144]	[0.098]	[0.153]	[0.064]	[0.146]	[0.216]	[0.247]
ln(health capital per worker)	0.005	0.143	0.061	0.252*	0.072	0.316**	0.485***	0.568**
	[0.100]	[0.118]	[0.069]	[0.130]	[0.087]	[0.145]	[0.168]	[0.186]
Linear time trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Structural break dummy	No	Yes	No	Yes	No	Yes	No	Yes
Prewhitened	N/A	N/A	Yes	Yes	Yes	Yes	No	No
Sum of squared residuals	0.115	0.104	0.118	0.114	0.117	0.123	0.037	0.033

Notes: Dependent variable: ln(GDP per worker), ln(•)=natural logarithm transformation. All models include constants. Besides OLS estimators, HAC standard error estimates are in brackets. *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A: Not applicable. The leads and lags are set to be 1 for DOLS estimators.

Table 2.8. DOLS Estimations of the long-run cointegrating relationship, 1963-2007.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
ln(physical capital per worker)	0.167	0.194	0.144	0.163	0.043	0.205	-0.013	0.172	0.010
	[0.219]	[0.154]	[0.179]	[0.185]	[0.168]	[0.203]	[0.168]	[0.233]	[0.248]
ln(education capital per worker)	0.586**	0.548***	0.638***	0.540**	0.725***	0.418*	0.716***	0.486*	0.623**
	[0.247]	[0.161]	[0.190]	[0.204]	[0.185]	[0.240]	[0.189]	[0.268]	[0.287]
ln(health capital per worker)	0.568**	0.514***	0.550***	0.522***	0.530***	0.397**	0.468***	0.402**	0.335
	[0.186]	[0.123]	[0.145]	[0.135]	[0.156]	[0.158]	[0.158]	[0.195]	[0.219]
Lags / Leads	1/1	0/0	0/1	1/0	0/2	2/0	1/2	2/1	2/2
Linear time trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Structural break dummy	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sum of squared residuals	0.033	0.039	0.036	0.035	0.027	0.031	0.023	0.028	0.021

Notes: Dependent variable: ln(GDP per worker), ln(•)=natural logarithm transformation. All models include constants. HAC standard error estimates are in brackets. *, **, and *** indicate the significance level at 10, 5, and 1% respectively.

Table 2.9. Estimations of the long-run cointegrating relationship, 1963-2007 (per capita form).

	OLS		FMLS		CCR		DOLS	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(physical capital per capita)	0.135	0.190	0.100	0.045	0.038	0.119	0.183	0.268
	[0.180]	[0.169]	[0.209]	[0.185]	[0.167]	[0.133]	[0.227]	[0.184]
ln(education capital per capita)	0.487***	0.354**	0.547***	0.469***	0.574***	0.448***	0.614***	0.487***
	[0.145]	[0.145]	[0.160]	[0.150]	[0.147]	[0.124]	[0.175]	[0.162]
ln(health capital per capita)	0.128	0.248***	0.256***	0.302***	0.225***	0.320***	0.424***	0.460***
	[0.083]	[0.090]	[0.093]	[0.094]	[0.080]	[0.090]	[0.135]	[0.106]
Linear time trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Structural break dummy	No	Yes	No	Yes	No	Yes	No	Yes
Prewhitened	N/A	N/A	No	Yes	No	Yes	No	No
Sum of squared residuals	0.127	0.108	0.136	0.123	0.132	0.125	0.028	0.026

Notes: Dependent variable: ln(GDP per capita), ln(•)=natural logarithm transformation. All models include constants. Besides OLS estimators, HAC standard error estimates are in brackets. *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A: Not applicable. The leads and lags are set to be 1 for DOLS estimators. I had difficulty in calculating prewhitening QS kernel estimators for Column (3) & (5), thus standard QS kernel estimators have been used.

To check for Granger causality for the cointegration relationship, a VEC model was estimated, the results are reported in Table 2.10. Starting with a weak exogeneity test, the theoretically predicted negative coefficients of ECT_{t-1} are only statistically significant in the GDP equation and the physical capital equation (Columns (1) and (2)). These results provide evidence regarding the debate on the health capital-economic output relationship, i.e. whether health capital positively causes GDP to grow, or the positive coefficient of health capital actually comes from GDP. The significant lagged error-correction term in the GDP equation (Column (1)), which is not seen in the health equation (Column (4)), implies that based on national time series data, in the long run, causality runs from health capital to GDP. As for the short-term effect, this study finds significant uni-directional Granger causality running from health capital and physical capital to GDP. There is evidence of bi-directional Granger causality between health capital and physical capital, education capital and physical capital, and between education capital and health capital.

Table 2.10. Results of Granger non-causality *F*-tests, 1963-2007.

	Δ ln(GDP per worker)	Δ ln(physical capital per worker)	Δ ln(education capital per worker)	Δ ln(health capital per worker)
	(1)	(2)	(3)	(4)
	Short-run Granger non-causality test			
Δ ln(GDP per worker)	–	1.59	0.77	0.01
Δ ln(physical capital per worker)	2.87*	–	3.02*	5.41**
Δ ln(education capital per worker)	2.09	4.32**	–	5.17**
Δ ln(health capital per worker)	3.85*	11.70***	7.41***	–
	Weak exogeneity test / long-run Granger non-causality test			
ECT_{t-1}	-0.87	-0.31	-0.16	-0.15
(*t*-statistic)	[0.17]***	[0.11]***	[0.13]	[0.17]

Notes: Δ is the first difference operator. *, **, and *** indicate the significance level at 10, 5, and 1% respectively. The lag selection is following Wooldridge (2009) with a maximum of 2 lags. *ECT* is estimated by FMLS in Column (4), Table 2.7.

Finally, over a shorter period (1978-2007), alternative data are available for measuring education capital stock, allowing some assessment of the robustness of the model estimates. The results are reported in Table 2.11. With regard to the key variable of interest, except in Column (5), all other efficient estimators show that health capital has a significantly positive role (the significant coefficients ranges from 0.451 to 0.732) at the 5% significance level in this period. However, contrary to the results reported using a longer time period (1963-2007), education capital (no matter which proxies are adopted) is insignificant while physical capital is now significant.

Table 2.11. Estimations of the long-run cointegrating relationship, 1978-2007.

	OLS		FMLS		CCR		DOLS	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(physical capital per worker)	0.625***	0.518***	0.400**	0.490***	0.545***	0.501***	0.191	0.564***
	[0.153]	[0.090]	[0.159]	[0.102]	[0.080]	[0.109]	[0.191]	[0.054]
ln(education capital per worker)	-0.086		0.051		-0.113		0.396	
	[0.162]		[0.169]		[0.099]		[0.282]	
ln(average years of schooling per worker)		-0.229		-0.072		-0.042		0.525
		[0.462]		[0.547]		[0.546]		[0.386]
ln(health capital per worker)	0.295*	0.332**	0.451**	0.471**	0.300	0.500**	0.732***	0.691**
	[0.153]	[0.156]	[0.169]	[0.181]	[0.186]	[0.203]	[0.221]	[0.130]
Linear time trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prewhitened	N/A	N/A	Yes	Yes	Yes	Yes	No	No
Sum of squared residuals	0.033	0.033	0.044	0.037	0.038	0.040	0.005	0.005

Notes: Dependent variable: ln(GDP per worker), ln(•)=natural logarithm transformation. All models include constants. Besides OLS estimators, HAC standard error estimates are in brackets. *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A: Not applicable. The leads and lags are set to be 1 for DOLS estimators.

To summarize, there is robust evidence suggesting health capital has exhibited a significantly positive effect on economic output in China while the effects of physical capital and education capital are mixed. Ideally, one would expect that both physical and education capital would have a significantly positive role in the production function. However, that is not always evidenced in the empirical analyses. For example, using China's provincial level data, Chi (2008) found that fixed assets investment is endogenous in economic growth regression, and after using instrumental variables to correct for it, the coefficient for fixed assets investment is insignificant. For education capital, several studies also found that education capital is insignificant in the growth model regardless of including health capital or not, see: Knowles and Owen (1997), McDonald and Roberts (2002), Bloom *et al*. (2004), McDonald and Roberts (2006), Bloom *et al*. (2010), Narayan *et al*. (2010), Audibert *et al*. (2012). Considering the limited time series, further robustness checks are undertaken in the next chapter when provincial-level data are adopted.

2.4. DISCUSSION AND CONCLUSIONS

As a cornerstone of human capital, improvements in health have been widely perceived as a major source of economic development. A national time series of 45 years has been utilised to investigate the effect of health investment on China's economic output. This chapter accounts for unknown structural breaks in the cointegrating relationship using both Gregory-Hansen and Hatemi-J approaches, and deals with possible endogeneity using

FMLS, CCR and DOLS estimation methods. A significant structural break is consistently found to occur in 1976, at the end of the Cultural Revolution (1966-1976). As anticipated, in the full sample models (which allow for the structural break and are estimated by three consistent estimators) health capital has a positive effect on GDP; the effect is still significant even when the sample size is reduced to the last 30 years. The robust positive effect of health capital shown here is consistent with both development growth theory and several empirical papers reviewed above, especially Li and Huang's panel data analysis on the Chinese economy. The conclusions for physical capital and education capital remain uncertain.

China's economy has undergone major changes since 1949; research into China's economic performance must utilize methods that allow for such changes. Several studies considering potential structural breaks in China's economy arbitrarily set 1978 as the break point, and then undertake modeling for the whole dataset and the two sub-samples separately. This approach works under the assumption that the exact timing of the break year is known. Both Gregory-Hansen and Hatemi-J approaches, as employed here, on the other hand, are residual-based tests that allow the data to determine the timing of structural break(s), so that they prevent informal data analysis from contaminating the choice of breakpoint(s) (Gregory and Hansen, 1996). It is more appropriate to let the data "tell" us the breakpoint and then the knowledge of the country's economic history can be used to explore the possible explanations for the existence of that break point.

As mentioned, the estimation suggests the structural break occurs in 1976, at the end of the Cultural Revolution. During this period, China suffered serious setbacks and losses (Du, 1992): from 1974 to 1976, financial revenue fell by 20,000 million Yuan; the national economy came to the very brink of collapse. China entered a new historical stage of development after the downfall of the "Gang of Four" in 1976. In 1977 the total value of industrial and agricultural output increased by 10.7% over that of 1976; revenue increased by 12.6% leading to a surplus of 3,100 million yuan (Du, 1992). As also noted by Bramall (2009), Mao’s death in 1976 inevitably led to the gradual unraveling of the development strategy which had been pursued since 1963. This Revolution also had a profound impact on China’s education system. During the revolution, the education system was completely revamped (World Bank, 1983). The universities were closed between 1966 and 1970 and operated in response to political instead of academic considerations from 1971 to 1976 (Guo, 2009). Since 1976, China’s educational policy has again been modified: to provide the manpower necessary for rapid economic development, the Government decided to give increased priority to education generally and in particular to concentrate additional resources on specialized institutions and higher education (World Bank, 1983). The revolution did not have a direct effect on the health system in China; however, during the years afterwards, the health care system transited through economic reform under hard budget constraints and local fee-generated revenues (see Chapter 1 for details). In summary, there are several good reasons as to why 1976 shows up as the significant structural break in the period under investigation.

Two main limitations to this study are summarised here as possible directions for future research. Firstly, the measurement of health capital chosen in this study is not ideal, limited mainly due to data availability. Life expectancy (the most popular proxy in this field) or disability-adjusted life years (DALYs)[17] may be a better choice; however, such annual data for China are unavailable. For cross-country analysis, the data for DALYs is also limited in that the available information can only be traced back to 1999 (see Audibert *et al*., 2012). Using life expectancy may also give rise to other issues concerning whether extending life years actually captures the health effect or the work experience effect. This makes definitively interpreting the growth effect of the health capital proxy difficult to explain in this case.

In the production function context, the health capital effect estimated is comparable to other results in the literature based on cross-country data and adopting life expectancy as a proxy. For example, implied elasticities of health capital reported by Knowles and Owen (1995, page 104) are 0.328 and 0.381 for all 84 countries and 0.382 for 62 less developed countries based on OLS estimates; implied elasticities of health capital reported by Knowles *et al.* (2002, page 141) range from 0.422 to 0.515 when the Two-stage Least Squares (2SLS) method has been used; by adopting panel analysis, implied elasticities of health capital reported by McDonald and Roberts (2002, page 274) are 0.18, 0.16 and 0.42

[17] DALYs are defined as "the sum of years of potential life lost due to premature mortality and the years of productive life lost due to disability" by WHO.
(see: http://www.who.int/healthinfo/global_burden_disease/metrics_daly/en/)

for all 77 countries, 55 less developed countries and 39 less developed countries (which is the former 55 countries less Latin American countries) respectively.

A second limitation is that the sample size of this study is relatively small, making long run effects difficult to estimate with precision. A panel cointegration analysis based on provincial data is a logical next step; the variation across provinces will give considerably more information with which to estimate the model.

Although health investment may significantly contribute to economic growth for less-developed countries, under-investment is quite common in those countries. One potential reason for this phenomenon may be what Chen (2010) describes as a traditional misconception in developing countries that spending on health is seen as a social burden, rather than a strategic investment vital for each nation's socioeconomic development. The break-out of Severe Acute Respiratory Syndrome (SARS) is a direct lesson that public health crises can produce enormous economic losses, affecting short-run economic growth (Wang, 2004). In contrast, investment in health may have a more positive long-run effect on a country’s economic growth. There was a period where the Chinese government was shifting the main burden for health care financing to its population; however, since the early 2000s, government health expenditure has risen again. Within China’s Health Care Reform Plan for 2009-2011, the Chinese government promised to spend 850 billion yuan (U.S. $125 billion) by 2011 to begin correcting the country's health system and it makes the first concrete step towards a goal of providing primary health services to all Chinese people by the year 2020 (Han *et al*., 2010). The government's increasing role in financing healthcare in the new Health Reform Plan is an important symbol of the government's health investment strategy and serves a key condition for a sustainable health system. This chapter's analysis suggests that beyond the direct health benefits, a health investment strategy can also contribute to sustainable economic growth in light of the role played by health human capital.

CHAPTER 3

HEALTH INVESTMENT AND ECONOMIC OUTPUT IN REGIONAL CHINA

3.1. INTRODUCTION

The relationship between health capital and economic output has been the subject of a great deal of theoretical and empirical investigation (see Chapter 2 for a review). While the positive effects of health on income and economic output are well established, the evidence is limited as whether or how this effect varies with the level of development of a country or region. Conclusively estimating the dynamics of the health-growth relationship via the economic history of just one country would require a long time series to allow for enough variability in levels of health capital and of aggregate economic activity. An alternative is to undertake a cross-country panel study. The diversity of levels of Gross Domestic Product (GDP) and health investment, and of their trajectories across time, creates the possibility of discerning the impact of health capital on growth at various levels of investment and of income. However, there are a number of criticisms of cross-country studies, not least that because of the wide degree of heterogeneity across countries, unified models are potentially implausible (Maddala, 1999).

The Chinese experience offers something close to a natural experiment that sheds light on this important question in a way that avoids some of the obvious pitfalls of cross-country studies. Specifically, China's recent history has seen a great deal of geographical diversity in the levels of investment in health, and in levels of economic growth, as a result of largely exogenous policy decisions of a Central government. The geographical diversity can be used to model the variations in how health investment affects economic activity. Therefore, regional economic activity in China is the focus of this chapter. A panel data set of Gross Regional Product (GRP) for all the provinces of China over a 29 year period was constructed. This is used to estimate the effect of health investment on growth in this output measure. By comparing model estimates for the different major regions of China (specifically, the coastal provinces and the inland provinces), insight is provided into how health investment can affect growth differently at different stages of development, as these two regions have developed very differently over the sample period.

Li and Huang's (2009) study provides the most closely related previous work on this topic. They use annual data from 28 provinces and an augmented Mankiw-Romer-Weil (MRW) framework, adopt panel data techniques (fixed-effects (FE) and random-effects (RE) regression analysis), and find that health capital (proxied by the number of doctors per

10,000 people or the number of hospital beds per 10,000 people) has a positive effect on GRP per capita. Their sensitivity analyses found that the health effect is stronger in coastal China compared to inland China. Given that the average number of doctors per 10,000 people was 15.29 in the coastal region and 13.92 in the inland region during the periods from 1978 to 2006[18], this finding from Li and Huang tends to suggest increasing returns to health capital. However, for these long panel datasets, whether variables are stationary or not needs careful examination so as to avoid spurious regression problems. Besides, even if the series are cointegrated, unlike in standard time-series analysis, the FE estimator is inconsistent in the panel data setting if the series are nonstationary (Baltagi, 2005).

The econometric challenges of a study such as this are not trivial. This chapter carefully addresses each of these challenges, utilising cutting edge econometric techniques, with the aim of providing robust conclusions about how the impacts of health investment on aggregate economic activity vary with the stages of development. The results presented here show that careful consideration of these econometric issues can make a large difference to the results and conclusions reached. For example, this chapter finds the effect of health investment is much stronger in inland China compared to the coastal region, a sign of diminishing returns to health capital.

To put this chapter in the context of earlier literature, firstly an overview of some of the key previous work, both theoretical and empirical, is presented. Knowles and Owen (1995) considered health capital in the classic MRW exogenous growth model; in this case health can be regarded as separate from physical capital that may influence the level of output (e.g. GDP). Based on the above framework, empirical evidence from Knowles and Owen

[18] In 2006, the number of doctors per 10,000 people in the municipality of Beijing is 33.39 (the highest in eastern regions), three times larger than Guizhou Province in which the figure is 10.40 (the lowest in inland regions).

(1995), (using data from 84 countries and adopting a transformation of life expectancy as a proxy for health capital) and McDonald and Roberts (2002) (using data from 77 countries and using infant mortality and a transformation of life expectancy as proxies for health capital, respectively), suggest that health capital is significant for less developed countries.

Using a production function framework, Benos and Karagiannis (2010) used regional panel data from Greece, and found that health capital (proxied by the number of medical doctors) is more significant for economic growth in poor than rich regions. Bhargava *et al*. (2001) includes both health status (adult survival rates, ASR) and an interaction term between health status and the level of GDP in the growth equation to allow for some form of non-linearity. The authors suggest that owing to the contribution of labour in prime years, for countries at low levels of GDP, adult survival rates have a significant positive effect on economic growth. Gyimah-Brempong and Wilson (2004) use data from Sub-Saharan Africa and Organisation for Economic Co-operation and Development (OECD) countries and estimate the growth equation separately for the two samples. By including variables measuring investment in health capital (the ratio of total health expenditure to GDP for 23 OECD countries and the ratio of government health expenditure to GDP for 21 Sub-Saharan African countries), health capital stock (the inverse of the child mortality rate per 1000 births) and its squared term, they found robust results from the two samples that suggest the growth effect of health capital stock is positive but subjected to diminishing returns. Comparing the coefficients obtained from the Sub-Saharan African sample and OECD countries sample, evidence suggests that health capital stock has a much larger impact on growth in less developed countries where the stock of health is lower.

Analyses of endogenous growth models provide a similar conclusion from a different perspective. Zhang and Zhang (2005) consider life expectancy in an endogenous growth framework and find that rising longevity will have a positive effect on growth at a diminishing rate. Leung and Wang (2010) further investigate the endogenous nexus of health care, life expectancy and economic output in a neoclassical growth model. Their model separates the health input (health care) from the health output (life expectancy) and shows that health care helps prolong life expectancy, leading to higher savings and hence capital formation; thus for those economies who spend more on health care there will be higher output in the long run. By assuming the survival rate is concave, the marginal benefit of health investment diminishes.

Further microeconometric evidence also suggests similar conclusions, that health human capital has a positive effect on income, but with diminishing returns. Schultz (2002) found that an increase in height has a positive effect on wages, but the effect is stronger in (the less developed nations of) Ghana and Brazil than in the United States. Liu *et al*. (2008), using survey data from China, find that household income is strongly influenced by self-reported health status, particular in rural areas.

China is a very large developing country with one of the most rapidly growing economies in the world, with a real average GDP per capita growth rate of 7.2 percent from 1978 to 2006. Whether health investment can help explain the economic growth experience of China has become a topic of increased interest in recent years. In this context, as noted, the Li and Huang (2009) paper is particularly relevant, because a province level dataset has been used. A more recent study by Gong *et al*. (2012) uses annual data from 28 provinces (over the period of 1978-2003) and found that the health capital growth rate (proxied by the number of hospital beds) and the level of health capital both significantly impact on economic growth in China. By using cross-country data, Bloom *et al*. (2010) show that improved health (as measured by life expectancy) can explain the growth in both China and India. However, also adopting cross-country data, Jamison *et al*. (2010) suggest that the effect of health on growth in China is close to zero.

It is hypothesized that investment in health should have a significant impact on economic output in China. A clear recent example is the outbreak of Severe Acute Respiratory Syndrome (SARS) in 2003, which drew unprecedented attention to the importance of health and health care, and China received a direct lesson that public health crises can induce enormous economic losses, affecting short-run economic growth (Wang, 2004).

Using 45 years of national time series data, the result based on a cointegration analysis in Chapter 2 of this thesis finds the robust conclusion that health capital has a significantly positive effect on economic output. This chapter builds upon this analysis with a rich provincial-level panel dataset.

The chapter is structured as follows: the empirical model and econometric method are presented in Section 3.2; followed by results in Section 3.3; finally, discussion and conclusions are summarised in Section 3.4.

3.2. METHODOLOGY AND DATA

3.2.1. Theoretical Model

Consistent with Chapter 2, a production function extended to include human capital in two forms – education and health – is considered and investigated in a panel data framework. The model assumes an aggregate production function with constant returns to scale (CRS)[19]. The Cobb-Douglas production function with factor-augmenting technological progress, for time period *t*, is:

$$Q_{it} = A_{it} F(K_{it}, E_{it}, H_{it}, L_{it}) = A_{it}^{\gamma} \cdot K_{it}^{\alpha} \cdot E_{it}^{\beta} \cdot H_{it}^{\eta} \cdot L_{it}^{\chi} \tag{3.1}$$

where $\chi = 1 - \alpha - \beta - \eta$. It is assumed that $\gamma, \alpha, \beta, \eta > 0$ and $0 < \chi < 1$. The notation adopted is as follows: *Q* is aggregate output, *K* is the stock of physical capital, *E* and *H* are variables that express the stock of human capital, education and health respectively, *A* is the level of technology and assumed to grow exogenously at rate *g* over time ($A_{it} = A_{i0} e^{gt}$), *L* is the level of employment, and the subscript *i*, *t* denotes the province and time period, respectively. Taking the natural logarithm of both sides of Equation (3.1) gives:

$$\ln Q_{it} = \alpha \ln K_{it} + \beta \ln E_{it} + \eta \ln H_{it} + \chi \ln L_{it} + \gamma \ln A_{i0} + \gamma g t \tag{3.2}$$

If *q* is defined as the level of output per worker, i.e., *q* = *Q* / *L*; and *k*, *e*, *h* as the stocks of physical capital, education capital stock and health capital stock per worker, respectively, Equation (3.2) can be further written as a per worker equation:

$$\ln q_{it} = \vartheta' x_{it} + \kappa t + \tau_i \tag{3.3}$$

[19] The CRS assumption is tested by two steps: firstly, the Equation (3.2) is estimated using a fixed-effects estimator, and secondly, an *F*-test is adopted to see if the null hypothesis ($\alpha + \beta + \eta + \chi = 1$) holds. The result suggests the CRS assumption cannot be rejected (*p* = 0.395).

where $x_{it} = \left(\ln k'_{it}, \ln e'_{it}, \ln h'_{it}\right)'$, $\kappa = \gamma g$, and $\tau_i = \gamma \ln A_{i0}$

3.2.2. Econometric Methodology

3.2.2.1. Panel unit root test

The first step of the analysis is to test whether the series are stationary or not. If the cross-sectional units (i.e. the provinces in this study) are not randomly sampled, then cross-section independence is a highly unrealistic assumption. For example, the regions are likely to be vulnerable to common policy shocks or business cycles. Thus, a cross-section correlation robust panel unit root test is a suitable choice. Among a group of candidate tests, the one proposed by Breitung and Das (2005) is chosen since it does not need to estimate the unobservable common factor representation[20]. As the (feasible) generalised least squares (GLS) *t*-statistic is used, the number of time periods (T) should be no less than the number of cross-section units (N). Considering this issue, the test is performed on the two samples (coastal and inland China) separately.

3.2.2.2. Panel cointegration tests

Following the panel unit root tests, if the series are non-stationary, a number of residual-based panel cointegration tests based on augmented Dickey-Fuller (ADF) regressions will be utilised to avoid spurious regressions. Firstly, two classical panel cointegration tests with a null of no cointegration, proposed by Pedroni (1999, 2004), are

[20] The test results may be subject to size distortions due to a small sample size; and secondly, the results may be biased since in the analysis the T is only relatively larger than N (Breitung and Das, 2008).

used. Pedroni (1999, 2004) developed seven tests[21] that belong to two classes of statistics: pooled within dimension tests and group mean between dimension tests. Pedroni found that the group-means estimators typically have lower small-sample size distortions than within-groups estimators. Among several single equation panel cointegration tests[22], Wagner and Hlouskova (2010) found that Pedroni's tests are those least affected by the presence of an $I(2)$ component (or two unit roots), short-run, cross-section correlation or cross-unit cointegration; specifically, two of the tests based on ADF regressions perform best. Secondly, the critical values for the Pedroni's group mean between dimension ADF statistics are calculated through the Continuous-Path Block bootstrap (CBB) method proposed by Fachin (2007). This bootstrap test is shown to have good power, especially when the sample size is small and more importantly is robust for both short- and long-run cross-section dependence. Finally, potential structural breaks are considered. Since the time period of data is relatively short (T = 29), only the case with one break in the level is considered, following Banerjee and Carrion-i-Silvestre (2006)'s two-step approach. This involves identifying the potential structural breaks for each region using Gregory and Hansen's (1996) approach, and then combining the region-specific *t*-ratio in a between-dimension test statistic as shown in Pedroni (1999, 2004). After properly standardising, the panel test statistics are shown to converge to a standard normal distribution (Banerjee and Carrion-i-Silvestre, 2006). Note that in this case (i.e. for a change in level), the power of the standard parametric Pedroni panel unit root tests will not be influenced by the misspecification error of the deterministic component (Banerjee and Carrion-i-Silvestre, 2006).

[21] Both the pooled within dimension tests and the group mean between dimension tests include the following three tests: a test analogous to the Phillips and Perron (PP) rho-statistic, a test analogous to the PP t-statistic, and a test similar to the ADF-type t-statistic; there is another non-parametric variance ratio statistic developed for the pooled within dimension tests.

[22] The single equation tests compared included: Pedroni (1999, 2004) and Westerlund (2005). Wagner and Hlouskova (2010) also compared two system panel cointegration tests of Larsson *et al*. (2001), as did Breitung (2005). However, they found that both those system tests show very bad performance for small values of T; they are also sensitive with respect to the presence of a $I(2)$ component.

3.2.2.3. Panel cointegration analyses

The fixed-effects within estimators are adopted first and the results are used as a reference point for the subsequent regression analyses. The cross-section dependent (CD) test proposed by Pesaran (2004) is applied to the FE estimates to see if cross-section dependence exists[23]. If cross-section dependence does exist after performing the tests mentioned above, a semi-parametric approach proposed by Driscoll and Kraay (1998) and coded by Hoechle (2007) will be adopted to calculate the heteroskedasticity, autocorrelation and cross-section dependence consistent standard errors.

As mentioned in the introduction section, the FE estimator is inconsistent in the cointegration analysis (Baltagi, 2005, pages 257-258). To deal with this issue (and also account for the endogeneity and serial correlation in the errors), a within-dimension panel dynamic ordinary least squares (DOLS) estimator, which adds the past and future values of Δx_{it} as additional regressors as proposed by Kao and Chiang (2000) and coded by Diallo (2010), will be adopted. The within-dimension panel DOLS estimator has been found to have good performance in Monte Carlo experiments in the one-dimensional cointegrating space, compared to several popular estimators[24] (Wagner and Hlouskova, 2010).

Including the leads and lags values of Δx_{it} as additional regressors, Equation (3.3) can be extended to:

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[23] A number of tests have been proposed to serve this aim. Moscone and Tosetti (2009) suggest this CD test performs the best.

[24] The estimators considered in their experiment included: panel fully modified (FM) OLS (Phillips and Moon, 1999), group-mean FMOLS (Pedroni, 2000), panel DOLS (Kao and Chiang, 2000), group-mean DOLS (Pedroni, 2001), one-step (group-mean) Vector autoregression (VAR) (Larsson and Lyhagen, 1999; Larsson *et al*., 2001), and two-step panel VAR (Breitung, 2005) estimators.

$$\ln q_{it} = \vartheta' x_{it} + \sum_{j=-q}^{q} c_{it}' \Delta x_{it+j} + \kappa t + \tau_i + v_{it}^* \tag{3.4}$$

where the specific form of error term v_{it}^* could be refer to Kao and Chiang (2000, 187-188). Heteroskedasticity and autocorrelation consistent (HAC) standard errors are calculated. To further (partially[25]) handle the short-term, cross-section dependence, common time dummies are included as suggested by Pedroni (2001). To account for potential structural breaks, the structural break dummies identified in the above panel cointegration test are also included.

When working with panel data, it is necessary to choose between the fundamental homogeneous estimator and the heterogeneous estimator. The homogeneous estimator is preferred in this chapter since the evidence from Baltagi *et al*. (2000) and Gutierrez *et al*. (2007) suggests that the pooling estimator is more economically plausible. The sub-sample (coastal and inland China samples) analyses may provide better estimators since they can be regarded as analyses on two sets of regions that are less heterogeneous, mitigating against possible heterogeneity of slope parameters. As reported by Keng (2006), the inter-regional disparity between inland and coastal China has been increasing since 1967, while the intra-regional disparity has decreased since then. The coastal-inland disparity accounted for 71% of overall regional disparity in China.

As robustness analysis, two heterogeneous estimators have also been used. The first is the Augmented Mean Group (AMG) estimator recently proposed by Eberhardt and Bond (2009), Eberhardt and Teal (2010) and coded by Eberhardt (2012). The AMG estimates the mean slope coefficient in a linear heterogeneous panel model and deals with the

[25] As discussed by Pedroni (2001, page 730), including common time dummies will not solve all kinds of cross-section dependence issues; for example, if dynamic feedback effects exist between the variables of different regions.

cross-section dependence by including a 'common dynamic process' in the cross-region regression. One drawback of this approach is that it does not handle the potential reverse causality among inputs and outputs in the production function. The last estimator adopted in this analysis is a between-dimension group-mean panel DOLS estimator proposed by Pedroni (2001). Compared with the homogeneous estimator, the two heterogeneous estimators allow for greater flexibility of the cointegrating vectors. However, when the time dimension is small, the group-mean panel DOLS could have a poor performance (Wagner and Hlouskova, 2010).

3.2.2.4. Panel Granger causality test

Using the production function framework and by adopting cointegration analysis, whether health capital has a significant positive effect on economic output can be tested. However, the direction of causality cannot be inferred from a cointegration equation. So as further verification, the final piece of analysis involves testing Granger causality in a panel context. The Engle and Granger (1987) two-step procedure is followed. Firstly, the estimated residuals of the cointegration analysis are saved. The one-period lagged residuals are then included into the following four dynamic panel equations:

$$\Delta \ln q_{it} = \theta_{1i} + \sum_{j=1}^{p} \theta_{11ip} \Delta \ln q_{i,t-j} + \sum_{j=1}^{p} \theta_{12ip} \Delta \ln k_{i,t-j} + \sum_{j=1}^{p} \theta_{13ip} \Delta \ln e_{i,t-j} + \sum_{j=1}^{p} \theta_{14ip} \Delta \ln h_{i,t-j} + \psi_{1i} ECT_{t-1} + u_{1it} \tag{3.5}$$

$$\Delta \ln k_{it} = \theta_{2i} + \sum_{j=1}^{p} \theta_{21ip} \Delta \ln k_{i,t-j} + \sum_{j=1}^{p} \theta_{22ip} \Delta \ln q_{i,t-j} + \sum_{j=1}^{p} \theta_{23ip} \Delta \ln e_{i,t-j} + \sum_{j=1}^{p} \theta_{24ip} \Delta \ln h_{i,t-j} + \psi_{2i} ECT_{t-1} + u_{2it} \tag{3.6}$$

$$\Delta \ln e_{it} = \theta_{3i} + \sum_{j=1}^{p} \theta_{31ip} \Delta \ln e_{i,t-j} + \sum_{j=1}^{p} \theta_{32ip} \Delta \ln q_{i,t-j} + \sum_{j=1}^{p} \theta_{33ip} \Delta \ln k_{i,t-j} + \sum_{j=1}^{p} \theta_{34ip} \Delta \ln h_{i,t-j} + \psi_{3i} ECT_{t-1} + u_{3it} \quad (3.7)$$

$$\Delta \ln h_{it} = \theta_{4i} + \sum_{j=1}^{p} \theta_{41ip} \Delta \ln h_{i,t-j} + \sum_{j=1}^{p} \theta_{42ip} \Delta \ln q_{i,t-j} + \sum_{j=1}^{p} \theta_{43ip} \Delta \ln k_{i,t-j} + \sum_{j=1}^{p} \theta_{44ip} \Delta \ln e_{i,t-j} + \psi_{4i} ECT_{t-1} + u_{4it} \quad (3.8)$$

Two types of causality tests are used, the weak exogeneity test (whether the coefficients of the one-period lagged error-correction term, ECT_{t-1}, derived from the long-run cointegration equation is significant or not) and the short-run Granger non-causality test (whether the coefficient of first differenced lagged variables are significant or not). A homogenous dynamic panel data estimator (two-step Generalised Method of Moments (GMM)) proposed by Arellano and Bover (1995) and Blundell and Bond (1998) has been adopted to handle the lagged dependent variable issue following Hartwig (2010). Stata Version 11.1 software, RATS Version 7.1 software, and GAUSS Version 9.0 software are used for empirical analyses.

3.2.3. Data and Variables

The dataset comprises 30 provinces of mainland China[26] (the municipality of Chongqing is merged into Sichuan Province) and 29 time periods (1978 to 2006). All data except real physical capital stock come from the *China Compendium of Statistics 1949-2004* and the *Provincial Statistical Yearbook*. Real Gross Regional Product (GRP) (constant 1978 price) is calculated by combining the nominal value of GRP in 1978, and the GRP price index.

[26] Matsuki and Usami (2011) argue that the difference between the Guangxi Zhuang autonomous region and other inland regions are less than those with coastal regions. Specifying the Guangxi Zhuang autonomous region as inland China, the main conclusions remain the same.

Real physical capital stock (constant 1978 price) comes from Wu (2009). A real physical capital stock variable is also available from Zhang (2008) and Wu (2008). The one reported in Wu (2009) is chosen here because it provides a longer period of data, since 1978. The three physical capital stock series are highly correlated, with the lowest simple correlation of 0.914. The number of employed persons is used as the level of employment.

Per worker education capital is defined as the number of students enrolled in secondary school per 10,000 employed persons. A similar proxy has been used by Chen and Fleisher (1996), Li *et al*. (1998), Wei *et al*. (2001) and Zheng *et al*. (2006). This variable may not be the ideal proxy for educational human capital stock as discussed by Chi (2008). However, it is the best available for this time period. As a robustness check, the average years of schooling per worker are predicted for the period 1978 to 2006 based on the available information from 1996 to 2006 in each province (see Appendix 3 for details). The proxy for health capital per worker is a health input measure, the number of doctors per 10,000 employed persons, as used and discussed in Chapter 2. In the robustness analysis, the number of hospital beds per 10,000 employed persons has also been adopted. See Table 3.1 for summary statistics of the variables.

Table 3.1. Summary statistics.

Variable		Mean	Std. Dev.	Min	Max	Obs.
ln(GRP per worker)	overall	7.774	0.846	6.092	10.730	870
	between		0.521	6.908	9.369	30
	within		0.673	6.352	9.355	29
ln(physical capital per worker)	overall	8.961	0.933	6.608	11.811	870
	between		0.572	8.219	10.327	30
	within		0.745	7.182	10.984	29
ln(education capital per worker)	overall	6.907	0.370	5.186	8.068	870
	between		0.283	5.835	7.342	30
	within		0.244	6.259	8.177	29
ln(health capital per worker)	overall	3.475	0.371	2.748	4.453	870
	between		0.362	2.990	4.322	30
	within		0.104	3.063	3.737	29

3.3. RESULTS

3.3.1. Panel Unit Root Test Results

The panel unit root test results are reported in Table 3.2 and the test results suggest that nonstationarity cannot be ruled out in this analysis and all series are integrated of order 1.

Table 3.2. Breitung and Das (2005) panel unit root tests results.

	with trend		without trend	
	lag = 0	lag = 1	lag = 0	lag = 1
Panel A: Coastal China				
ln(GRP per worker)	1.114	-0.769	7.869	2.681
Δ ln(GRP per worker)			-4.016***	-3.346***
ln(physical capital per worker)	2.994	0.097	7.608	1.820
Δ ln(physical capital per worker)			-3.739***	-2.452***
ln(education capital per worker)	1.490	-0.457	0.378	-1.248
Δ ln(education capital per worker)			-1.361*	-1.235
ln(health capital per worker)	0.864	0.069	-0.874	-1.120
Δ ln(health capital per worker)			-5.779***	-4.066***
Panel B: Inland China				
ln(GRP per worker)	2.683	0.714	9.948	4.864
Δ ln(GRP per worker)			-4.059***	-2.625***
ln(physical capital per worker)	4.604	0.929	11.052	3.280
Δ ln(physical capital per worker)			-3.527***	-0.407
ln(education capital per worker)	2.660	0.314	0.470	-1.215
Δ ln(education capital per worker)			-2.494***	-1.366*
ln(health capital per worker)	0.528	0.002	-1.019	-1.310*
Δ ln(health capital per worker)			-6.358***	-3.335***

Notes: Null hypothesis is that all series are non-stationary. Δ is the first difference operator. *, **, and *** indicate the significance level at 10, 5, and 1% respectively.

3.3.2. Panel Cointegration Results

The panel cointegration test results are reported in Table 3.3. The results suggest that there exists a cointegration relationship, allowing for unknown structural breaks or not. When the critical values are calculated through a CBB, which has more power when the sample size is small and when cross-section dependence may be an issue, this confirms a cointegration relationship. All robustness analysis results reported below passed the Pedroni cointegration test.

Table 3.3. Residual-based panel cointegration test results.

	Test	Statistics
Panel A: Pedroni (1999, 2004) (individual intercept and individual trend)		
Group-means between-dimension		
	ADF test	-2.780***
Pooled within-dimension		
	ADF test (weighted)	-2.156**
Panel B: Fachin (2007) (individual intercept and individual trend)		
Group-means between-dimension		
	ADF test	-3.611**
Panel C: Banerjee and Carrion-i-Silvestre (2006) (level shift with trend model, one break case)		
Group-means between-dimension		
	ADF test	-28.986***

Notes: Tests results reported from Panel A and C are calculated based on cross-sectionally demeaned data. For Panel B, the bootstrap is based on 1,000 redrawings, and the block length is 3. Regardless of whether simple bootstrap or fast double bootstrap critical values are used, the test-statistic is significant at 5% criteria. *, **, and *** indicate the significance level at 10, 5, and 1% respectively.

The estimated dates of structural breaks for each province are reported in Table 3.4. There is no clear pattern for the break dates. For the sake of partly controlling for cross-sectional dependence, the break dates are estimated based on cross-section demeaned data. For this reason it is difficult to tease out the potential reasons behind the identified breaks in a system of variables. Furthermore, as post-reform data have been used in this study, no major episode, such as the Cultural Revolution, which has severely affected the whole economy can be identified (see discussion in Chapter 2 for details). Thus the effect of structural breaks on the cointegration vector might be minor. This is supported in the subsequent regression results; see Columns (4) and (5) in Table 3.6 for example.

Table 3.4. Estimated structural breaks of each region, 1978-2006.

Region	Province	Year
Coastal	Beijing	1999
	Tianjin	1992
	Hebei	1983
	Liaoning	1991
	Shanghai	2000
	Jiangsu	1983
	Zhejiang	1991
	Fujian	1991
	Shandong	1998
	Guangdong	1989
	Guangxi	1991
	Hainan	2000
Inland	Shanxi	1987
	Inner Mongolia	1991
	Jilin	1998
	Heilongjiang	2002
	Anhui	1988
	Jiangxi	1987
	Henan	1999
	Hubei	1983
	Hunan	1998
	Sichuan & Chongqing	1997
	Guizhou	1983
	Yunnan	1989
	Tibet	1986
	Shaanxi	1988
	Gansu	1987
	Qinghai	1991
	Ningxia	1992
	Xinjiang	1993

Note: The breaks are assumed to be in the levels and identified based on cross-sectionally demeaned data.

Before presenting the cointegration results, simple correlation analysis of variables used in the key cointegration equation (Table 3.6) is shown in Table 3.5. All four series are positively correlated. Similar to those reported in Chapter 2, the physical capital and GRP series are especially highly correlated.

Table 3.5. Simple correlation statistics.

	ln(GRP per worker)	ln(physical capital per worker)	ln(education capital per worker)	ln(health capital per worker)
ln(GRP per worker)	1			
ln(physical capital per worker)	0.930	1		
ln(education capital per worker)	0.134	0.191	1	
ln(health capital per worker)	0.433	0.439	0.128	1

The regression results for the whole of China (30 provinces) are presented in Table 3.6. The first five columns report the results based on homogeneous estimators while the last two columns rely on heterogeneous estimators. The first two columns report the results from FE estimators. Time effects are allowed for by including a linear time trend or common time dummies in each year. As can be seen, coefficients of the three key variables are all significant and have the anticipated positive sign. However, since the FE estimates are inconsistent, these results should be viewed with some caution. The issue regarding cross-section correlation can be tested on the residuals of the FE estimates. The Pesaran's panel CD test after the FE estimate in Column (1) (in Table 3.6) suggests that cross-section correlation is significant (test statistics = 21.595, p = 0.000); after including the common time dummies, the CD test statistic suggests that the cross-section correlation is insignificant (test statistic = -0.664, p = 0.507). This provides some evidence that adding common time dummies makes the cross-section correlation issue less of a worry in this study. The following discussion is based on the results from the different efficient estimators (i.e. DOLS and AMG).

The within-dimension panel DOLS estimates are reported in Columns (3) to (5). The time effect is captured as a linear time trend in Column (3), or as common time dummies in Column (4). The magnitudes of the cointegrating vector for the three key variables are almost identical. Since using common time dummies can help in capturing unobservable common shocks, this method is adopted. The results from the homogeneous panel DOLS estimator (Columns (3) to (5)) found a significantly positive effect of physical and health capital. However, contrary to expectations, education capital is found to be insignificant or significantly negative. This result still holds when structural break dummies for each province are included, as reported in Column (5). Whether structural break dummies are included or not, the magnitude of the cointegration vector varies little. This may be because there was no major episode (such as the Cultural Revolution, which has severely affected the whole economy) during the post-reform time period, or that any structural change has not disturbed the relationship between health and GRP.

As discussed in Chapter 2, changing the number of leads and lags of the first difference control variables may have an impact on the estimated long-run coefficient. To see if this is the reason behind the unexpected results for education capital and also to test for the robustness of results reported in Columns (3) to (5) of Table 3.6, a set of different combinations of leads and lags are specified; the results are reported in Panel A in Table 3.7. The results are quite robust, in that there are positive coefficients on physical and health capital while the coefficients of education capital are all negative and mostly significant.

Panel B of Table 3.7 presents robustness results for excluding variables or some provinces. The coefficient of education capital remains negative when excluding health capital

(Column (8)), and tends to be positive when further excluding physical capital (Column (8)). Columns (10) to (12) report the robustness results for physical and health capital, which are found to be consistently significant and positive. To test for potential outlier effects the data from Tibet (Column (13)), and then both Tibet and Beijing (Column (14)) are excluded. However, the results are comparable with the key findings so far: the effects of both physical and health capital are significantly positive, but not for education capital.

Finally, the results based on two heterogeneous estimators are reported in Columns (6) and (7) in Table 3.6. Estimated long-run coefficients of physical, education and health capitals are all significant and positive. To summarise, for the whole China sample, there is robust evidence suggesting the long-run cointegrating coefficients of physical and health capital are significantly positive, while there is some evidence suggesting education capital is also significantly positive, although this is not as robust as the results for physical and health capital.

Table 3.6. Estimates of the long-run cointegrating relationship (whole China).

	Homogeneous					Heterogeneous	
	FE	FE	DOLS-WD	DOLS-WD	DOLS-WD	AMG	DOLS-GM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ln(physical capital per worker)	0.223***	0.194***	0.599***	0.585***	0.583***	0.412***	0.338***
	[7.15]	[7.27]	[11.31]	[11.33]	[11.29]	[6.02]	[33.28]
ln(education capital per worker)	0.225***	0.106**	-0.015	-0.084**	-0.090**	0.104***	0.079***
	[5.94]	[2.12]	[-0.40]	[-1.79]	[-1.92]	[3.53]	[7.88]
ln(health capital per worker)	0.283***	0.552***	0.283***	0.318***	0.311***	0.091**	0.128***
	[3.83]	[6.80]	[3.72]	[3.63]	[3.56]	[1.99]	[9.82]
Lags & Leads	N/A	N/A	1	1	1	N/A	1
Linear time trend	Yes	No	Yes	No	No	Yes	No
Common Time Dummies	No	Yes	No	Yes	Yes	Yes	Yes
Structural Breaks	No	No	No	No	Yes	No	No
Obs.	870	870	780	780	780	870	780
No. of Provinces	30	30	30	30	30	30	30

Notes: Dependent variable: ln(GRP per worker), ln(•)=natural logarithm transformation. The t-statistics are reported in brackets. FE is fixed-effects estimator; AMG is the augmented mean group estimator proposed by Eberhardt and Bond (2009), Eberhardt and Teal (2010); DOLS-WD is a within-dimension panel dynamic OLS estimator proposed by Kao and Chiang (2000); DOLS-GM is a group-mean panel dynamic OLS estimator proposed by Pedroni (2001). *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A - Not applicable.

Table 3.7. Estimates of the long-run cointegrating relationship (whole China).

Panel A: Robust analyses I (changing leads and lags)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ln(physical capital per worker)	0.572***	0.573***	0.575***	0.595***	0.587***	0.604***	0.605***
ln(education capital per worker)	-0.100**	-0.087**	-0.092**	-0.068*	-0.077*	-0.070*	-0.063
ln(health capital per worker)	0.348***	0.346***	0.336***	0.305***	0.305***	0.281***	0.268***
Lags / Leads	0/1	1/0	0/2	2/0	1/2	2/1	2/2
Obs.	810	810	780	780	750	750	720
No. of Provinces	30	30	30	30	30	30	30
Panel B: Robust analyses II							
	(8)	(9)	(10)	(11)	(12)	(13)	(14)
ln(physical capital per worker)	0.718***		0.578***		0.724***	0.583***	0.583***
ln(education capital per worker)	-0.063	0.178***				-0.113***	-0.112**
ln(health capital per worker)			0.340***	1.026***		0.334***	0.344***
Lags & Leads	1	1	1	1	1	1	1
Obs.	780	780	780	780	780	754	728
No. of Provinces	30	30	30	30	30	29	28

Notes: Dependent variable: ln(GRP per worker), ln(•)=natural logarithm transformation. The within-dimension panel dynamic OLS estimator proposed by Kao and Chiang (2000) is adopted. *, **, and *** indicate the significance level at 10, 5, and 1% respectively. All regression includes common time dummies. Column (13) excludes the data of Tibet; Column (14) further excludes the data of Beijing.

To check for Granger causality in the key cointegration relationship, two-step GMM estimators are adopted for Equations (3.5) to (3.8). Standard errors are estimated based on Windmeijer (2005). All models passed the Sargan overidentifying restriction tests and Arellano-Bond serial correlation tests (i.e. the null hypothesis of no serial correlations are expected to be rejected at order 1, but not at higher orders).

The theoretically predicted negative coefficients of ECT_{t-1} are statistically significant for all four equations, as reported in Table 3.8. The significant lagged error-correction terms in all equations imply that in the long run there exists causality in all directions among those four series. As for the short-term effect, this study finds significant uni-directional Granger causality running from physical capital to GRP, health capital to GRP, physical capital to education capital; bi-directional Granger causality between education capital and GRP, physical capital and health capital, and between education capital and health capital. See Table 3.8 for details.

Table 3.8. Panel Granger non-causality F-tests results.

	Δ ln(GRP per worker)	Δ ln(physical capital per worker)	Δ ln(education capital per worker)	Δ ln(health capital per worker)
	(1)	(2)	(3)	(4)
		Short-run Granger non-causality test		
Δ ln(GRP per worker)	–	2.42	5.98**	0.36
Δ ln(physical capital per worker)	34.88***	–	13.08***	5.79**
Δ ln(education capital per worker)	3.90**	0.12	–	4.04**
Δ ln(health capital per worker)	14.76***	16.41***	15.30***	–
		Weak exogeneity test / long-run Granger non-causality test		
ECT_{t-1}	-0.10***	-0.11***	-0.29***	-0.73***
(*t*-statistic)	[-2.94]	[-3.49]	[-8.79]	[-9.97]

Notes: Δ is the first difference operator. *, **, and *** indicate the significance level at 10, 5, and 1% respectively.

In the second part of this analysis, the regressions are run for the sub-samples separately. The results are reported in Table 3.9 for coastal China and Table 3.10 for inland China. Since the results based on FE estimators are inconsistent as discussed in the introduction, and the AMG estimates may confuse the reverse effect, as such only the results from two panel DOLS estimators are discussed here.

The estimates based on homogeneous estimators for the coastal China sample are similar to the full sample, in that physical and health capital are significantly positive while education capital is significantly negative. The results from the heterogeneous estimator find a slightly different result; education capital is significantly positive while health capital is now insignificant (physical capital is robustly significantly positive). The results from the inland China sample show robust, significant and positive coefficients for health capital. The coefficients on physical capital are all positive, but only significant in the heterogeneous estimates. Similar results have also been reported in Bloom *et al*. (2004) and as the authors discussed, this could be due to the lack of degrees of freedom (which is a common problem in macroeconomic studies). Education capital on the other hand, is found to be significantly positive in estimates that assume homogeneity but insignificant in heterogeneous panel DOLS estimates.

Next whether these estimates are significantly different between coastal and inland China are tested. For the preferred homogeneous panel DOLS estimates (Column (4) in Tables 3.9 and 3.10), the difference in physical capital effects is statistically significant (test statistic = 3.071, $p < 0.05$) while the difference for health capital is insignificant (test statistics = 0.615, $p > 0.10$). For the heterogeneous panel DOLS estimates (Column (7) in Tables 3.9 and 3.10), both physical capital and health capital show statistically significant differences in effect between coastal and inland provinces – the test statistic is 32.148 for physical capital ($p < 0.05$) and 2.134 for health capital ($p < 0.05$).

To briefly summarise the main findings, three main patterns can be observed. Firstly, health capital exhibits a positive effect on GRP, with significantly stronger effects in the inland China provinces than in the more developed coastal provinces (if heterogeneous estimates are adopted). Secondly, physical capital is found to have a positive effect on GRP, while the effect appears to be significantly stronger in coastal China using one set of estimates and stronger in inland China with other estimates. Thirdly, the effect of education capital is mixed.

Table 3.9. Estimates of the long-run cointegrating relationship (coastal China).

	Homogeneous					Heterogeneous	
	FE	FE	DOLS-WD	DOLS-WD	DOLS-WD	AMG	DOLS-GM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ln(physical capital per worker)	0.132***	0.121***	0.476***	0.438***	0.441***	0.449***	0.062***
	[4.29]	[4.20]	[6.92]	[6.70]	[6.79]	[3.55]	[18.73]
ln(education capital per worker)	0.292***	0.270**	0.011	-0.227***	-0.224***	0.069	0.179***
	[7.04]	[2.57]	[0.23]	[-3.03]	[-3.01]	[1.23]	[2.62]
ln(health capital per worker)	0.153*	0.288***	0.487***	0.546***	0.552***	0.030	-0.047
	[1.97]	[3.49]	[4.64]	[4.31]	[4.40]	[0.34]	[0.92]
Lags & Leads	N/A	N/A	1	1	1	N/A	1
Linear time trend	Yes	No	Yes	No	No	Yes	No
Common Time Dummies	No	Yes	No	Yes	Yes	Yes	Yes
Structural Breaks	No	No	No	No	Yes	No	No
Obs.	348	348	312	312	312	348	312
No. of Provinces	12	12	12	12	12	12	12

Notes: Dependent variable: ln(GRP per worker), ln(•)=natural logarithm transformation. The t-statistics are reported in brackets. FE is fixed-effects estimator; AMG is the augmented mean group estimator proposed by Eberhardt and Bond (2009), Eberhardt and Teal (2010); DOLS-WD is a within-dimension panel dynamic OLS estimator proposed by Kao and Chiang (2000); DOLS-GM is a group-mean panel dynamic OLS estimator proposed by Pedroni (2001). *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A - Not applicable.

Table 3.10. Estimates of the long-run cointegrating relationship (inland China).

	Homogeneous					Heterogeneous	
	FE	FE	DOLS-WD	DOLS-WD	DOLS-WD	AMG	DOLS-GM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ln(physical capital per worker)	0.228***	0.213***	0.052	0.029	0.039	0.348***	0.148***
	[6.77]	[7.21]	[0.97]	[0.66]	[0.90]	[4.48]	[15.75]
ln(education capital per worker)	0.166***	-0.024	0.247***	0.189***	0.174***	0.081**	-0.095
	[4.51]	[-1.30]	[6.85]	[4.86]	[4.48]	[2.09]	[0.65]
ln(health capital per worker)	0.208***	0.521***	0.588***	0.657***	0.647***	0.190***	0.411***
	[4.85]	[9.22]	[8.30]	[10.09]	[9.95]	[3.47]	[12.06]
Lags & Leads	N/A	N/A	1	1	1	N/A	1
Linear time trend	Yes	No	Yes	No	No	Yes	No
Common Time Dummies	No	Yes	No	Yes	Yes	Yes	Yes
Structural Breaks	No	No	No	No	Yes	No	No
Obs.	522	522	468	468	468	522	468
No. of Provinces	18	18	18	18	18	18	18

Notes: Dependent variable: ln(GRP per worker), ln(•)=natural logarithm transformation. The t-statistics are reported in brackets. FE is fixed-effects estimator; AMG is the augmented mean group estimator proposed by Eberhardt and Bond (2009), Eberhardt and Teal (2010); DOLS-WD is a within-dimension panel dynamic OLS estimator proposed by Kao and Chiang (2000); DOLS-GM is a group-mean panel dynamic OLS estimator proposed by Pedroni (2001). *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A - Not applicable.

Two further robustness analyses are undertaken as follows. Firstly, a complete series for the average years of schooling per worker for each province has been produced, based on interpolating available data, and used as an alternative proxy for education capital (see Appendix 3 for details). Secondly, the number of hospital beds was used as an alternative proxy for health capital. See Table 3.11-3.12 for detailed results. The coefficients on physical capital are significantly positive in all but one column. The results for education capital are found to be significantly positive in all samples when the average years of schooling proxy is adopted (see Table 3.11). This result suggests that education capital is not robust to different proxies. This issue is discussed in more detail in the following section. Finally, the health capital proxies have some impact on the magnitude of long-run elasticity, but a significantly positive effect can be found for both panel DOLS estimates (see Table 3.12).

Table 3.11. Estimates of the long-run cointegrating relationship: using average years of schooling as the proxy for education capital.

	whole China			coastal China			inland China		
	DOLS-WD	AMG	DOLS-GM	DOLS-WD	AMG	DOLS-GM	DOLS-WD	AMG	DOLS-GM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
ln(physical capital per worker)	0.352***	0.434***	0.339***	0.387***	0.353***	0.243***	0.022	0.430***	0.200***
	[6.94]	[7.94]	[27.05]	[6.19]	[4.40]	[17.42]	[0.53]	[6.532]	[12.14]
ln(average years of schooling per worker)	1.736***	0.343**	0.440***	4.622***	0.822***	1.526***	1.071***	0.239**	0.434***
	[7.20]	[2.54]	[5.45]	[11.59]	[2.69]	[9.33]	[6.60]	[2.083]	[3.86]
ln(health capital per worker)	0.179**	0.080*	0.365***	-0.480***	0.037	-0.103***	0.528***	0.190***	0.443***
	[2.05]	[1.71]	[6.74]	[-3.96]	[0.38]	[-5.43]	[8.23]	[3.721]	[15.40]
Lags & Leads	1	N/A	1	1	N/A	1	1	N/A	1
Obs.	780	870	780	312	348	312	468	522	468
No. of Provinces	30	30	30	12	12	12	18	18	18

Notes: Dependent variable: ln(GRP per worker), ln(•)=natural logarithm transformation. The t-statistics are reported in brackets. AMG is the augmented mean group estimator proposed by Eberhardt and Bond (2009), Eberhardt and Teal (2010); DOLS-WD is a within-dimension panel dynamic OLS estimator proposed by Kao and Chiang (2000); DOLS-GM is a group-mean panel dynamic OLS estimator proposed by Pedroni (2001). *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A - Not applicable. All regressions include time effect.

Table 3.12. Estimates of the long-run cointegrating relationship: using hospital beds as the proxy for health capital.

	whole China			coastal China			inland China		
	DOLS-WD	AMG	DOLS-GM	DOLS-WD	AMG	DOLS-GM	DOLS-WD	AMG	DOLS-GM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
ln(physical capital per worker)	0.642***	0.370***	0.227***	0.507***	0.370***	0.081***	0.253***	0.313***	0.297***
	[10.75]	[4.66]	[23.20]	[6.89]	[2.82]	[19.88]	[4.93]	[3.85]	[20.42]
ln(education capital per worker)	-0.160***	0.088***	0.199***	-0.406***	0.022	0.257***	0.032	0.047**	-0.171
	[-2.97]	[3.13]	[14.16]	[-5.26]	[0.36]	[5.23]	[0.68]	[1.96]	[1.25]
ln(hospital beds per worker)	0.291***	0.190***	0.149***	0.469***	0.219**	0.188***	0.478***	0.169	0.129***
	[3.36]	[2.78]	[11.96]	[3.85]	[2.06]	[7.54]	[7.16]	[1.26]	[5.86]
Lags & Leads	1	N/A	1	1	N/A	1	1	N/A	1
Obs.	728	812	728	312	348	312	416	464	416
No. of Provinces	28	28	28	12	12	12	16	16	16

Notes: Dependent variable: ln(GRP per worker), ln(•)=natural logarithm transformation. The t-statistics are reported in brackets. AMG is the augmented mean group estimator proposed by Eberhardt and Bond (2009), Eberhardt and Teal (2010); DOLS-WD is a within-dimension panel dynamic OLS estimator proposed by Kao and Chiang (2000); DOLS-GM is a group-mean panel dynamic OLS estimator proposed by Pedroni (2001). *, **, and *** indicate the significance level at 10, 5, and 1% respectively. N/A - Not applicable. All regressions include time effect.

3.4. DISCUSSION AND CONCLUSIONS

This chapter has investigated the role of health investment in explaining the level of GRP for mainland China, using 29 years of provincial-level panel data. Based on panel unit root tests and panel cointegration tests, the results suggest that there exists a cointegrating relationship among the variables. Taking account of the potential endogeneity of the independent variables, the homogeneous within-dimension panel DOLS estimator is adopted. The heterogeneous group-mean panel DOLS estimator and a recently proposed AMG estimator are also used as robustness checks. The main results suggest that health capital and physical capital are both vital for economic output in China; the effect of education capital is mixed. The magnitude of the coefficient on health capital is noticeably larger for less developed regions (inland China) than more developed provinces (coastal China) and is significantly larger based on heterogeneous estimates. This finding supports other empirical macro findings (as reviewed in the introduction section) using cross-country or cross-regional data that health capital has a greater impact on economies which are in less developed stages, as well as the empirical micro evidence that health status has a relatively stronger impact on personal wages (or household income) in developing countries (or regions), and also the theoretical conclusion that the marginal benefits of health investment are diminishing.

As for physical capital, although significantly positive effects have been found, the magnitude is sensitive to the estimation methods used. With education capital, the conclusions are mixed. The change in sign on education capital in explaining GRP can occur for a variety of reasons. Possibly the most fundamental reason is because there is not a good proxy for education capital *per se*. The number of students enrolled in secondary school is a raw measure of education capital and could be more akin to a flow measure instead of a stock measure. This proxy also clearly ignores the quality of education received. However, even using a more stock-related measure, such as average years of schooling per worker, there is not always a positive effect found in cross-sectional studies. For example, Bloom *et al*. (2004) separated the effect of schooling years and working experience years in a production function framework and found that, assuming common long-run total factor productivity (TFP) across countries, the effect of schooling years is positive but insignificant, while assuming country-specific long-run TFP, the effect of schooling changed to be negative and insignificant. These differ from results reported in other microeconomic studies, the authors concluding that they find no evidence of macroeconomic effects of education on economic growth and this may suggest the absence of externalities at the aggregate level. Using an updated dataset and a different model setting, Bloom *et al*. (2010) also report a negative coefficient of average schooling years on economic growth. Another potential reason the authors touch on, also supported by this study, is that education capital may impact on economic growth indirectly through its effect on health capital.

The study by Knowles *et al*. (2002) may provide a third explanation, that there may exist a gender difference effect of education capital on economic growth. The authors separate female and male education and study their effects separately. They found female education has a statistically significant positive effect on labour productivity, while the role of male education is not so clear. Therefore in a study using an aggregated proxy for education capital, as is the case here, the effect could potentially be less clear.

Several China-specific studies have examined the effects of education capital on growth.

Fleisher and Chen (1997), and Li and Huang (2009) found positive effects of education capital on economic growth, while Chen and Fleisher (1996), Li *et al*. (1998), Wei *et al*. (2001) found either insignificant or significant negative effects of education capital on GDP growth. Based on a shorter panel (1997-2006) and richer information, Zhang and Zhuang (2011) found that more developed provinces benefit more from tertiary education while less developed provinces depend on primary and secondary education. The limited data for the time period of this study do not allow me to explore these disaggregated effects.

For the key variable of interest in this chapter, health capital, the evidence suggests that there are diminishing returns. However, this conclusion should still be viewed with caution owing to the possibility of reverse causality. In panel cointegration analyses, the calculated long-run elasticity of the right-hand-side variables is the long-run cumulative multiplier only when the regressors are strictly exogenous. Results of this chapter show bi-directional Granger causality between health capital and GRP, suggesting the long-run multiplier cannot be consistently estimated. However, it is argued below that it can be reasonably assumed that any positive causal link from GRP to health capital will be of a similar magnitude or stronger for coastal China compared to inland provinces. This suggests that any bias in estimation of the long run multiplier for the effect of health on GRP will overstate the impact of health on growth by as much or more for coastal than inland provinces. It was found that the long run multiplier was much smaller for coastal provinces than inland, so the bias in estimates, if anything, leads to an underestimation of the gap between the two regions. In other words, when this potential bias is taken into account, the conclusion of diminishing returns to health is reinforced.

The argument concerning direction and strength of bias is built on the assumption that GRP has a stronger impact on health capital for the coastal provinces than for the inland provinces. The validity of this assumption is explored further here.

Following economic reform, the Chinese economy has experienced spectacular growth in the last three decades. However, during the decentralisation process, the central government experienced a drastic reduction in its revenue, which in turn reduced its capacity to fund health care (Yip and Hsiao, 2008), consequently responsibility for funding health services was transferred to local governments (Ho and Gostin, 2009). Under this situation, as noted by Huang and Luo (2009, page 203), “local governments in the poorest parts of China, which face the toughest public health challenges, spend the least on public health.” Although the disparity across the regions has been decreasing, the per capita government spending on health and education is still 1.5 times higher in coastal areas compared to inland China in 2006 (Huang and Luo, 2009).

The issue of under-investment in health care, especially in inland China, has been gradually recognised by the central government. Regional development policies – “Open up the West” or “Go West” (in 1999) and “Revive the Northeast” (in 2003) have been formally documented in the State Council Policy Document 2004 No. 6 (State Council Document No. 6, 2004). Enhancing health investment in these regions is one of ten issues highlighted by the document. Take Qinghai Province for example, “Go West” was implemented around a decade ago and the accumulated government investment in health care has reached 5.6 billion Yuan, exceeding the total amount of five decades’ of government health investment since the founding of new China; furthermore, the percentage of government health expenditure to total government expenditure increased from 1.9% in 1999 to 6.8% in 2009 (Health Department of Qinhai Province, 2010).

Recently, another regional development policy -"Central China Rising" (in 2004), further documented by the National Development and Reform Commission (National Development and Reform Commission Document No. 1827, 2010) also recognises enhancing health care reform and the public health system as key tasks that need to be implemented by 2015. Considering the transition of health investment strategies adopted by the central government, it is reasonable to conclude that the feedback effect of GRP on health investment was stronger in coastal China than inland China before 2000. It is this different feedback that more recent policy interventions were designed to address. It is thus reasonable to conclude that the result of diminishing returns to investment in health is robust to possible effects of reverse causality.

Two caveats are worth mentioning that may warrant future investigation when appropriate data becomes available. First, the measurement of health capital chosen in this study is not ideal, as discussed in Chapter 2. Health outcome variables would be a better choice; however, annual data for the provinces of China are unavailable at present. Reflecting on the results reported in Chapter 2, it can be seen that in the production function context, the magnitude of the health capital effect found in this thesis is comparable to other results in the literature where they adopt health outcomes (i.e. life expectancy and male survival rate) as proxies. See Table 3.13 for details.

Table 3.13. Comparison of implied health capital elasticities.

Sources	Method	Sample/Time period	Health capital/rate of health accumulation	Implied elasticities
Panel A: Cross country				
Knowles and Owen (1995)	OLS	84 countries / 1960-1985	-ln(80-life expectancy)	0.328 − 0.382
Knowles *et al.* (2002)	2SLS	73 countries / 1960-1990	-ln(85-life expectancy)	0.422 − 0.515
McDonald and Roberts (2002)	FE	77 countries / 1960-1989	-ln(80-life expectancy)	0.18 − 0.42
Jamison *et al.* (2010)	HLM	53 countries / 1965-1989	ln(male survival rate)	0.38 − 0.50
Rivera and Currais (1999)	OLS	24 OECD countries / 1960-1990	ln(health expenditure / GDP)	0.15 − 0.19
Narayan *et al.* (2010)	Panel DOLS	5 Asian countries / 1974-2997	ln(health expenditure / GDP)	0.168 − 0.268
Panel B: China				
Li and Huang (2009)	2SLS	28 provinces / 1978-2005	ln(no. of doctors per 10,000 people) / ln(no. of hospital beds per 10,000 people)	0.18 − 0.55
Chapter 2	FMLS / CCR / DOLS	national / 1963-2007	ln(no. of doctors per 10,000 workers)	0.252 − 0.586
Chapter 3	Panel DOLS	30 provinces / 1978-2006	ln(no. of doctors per 10,000 workers)	0.128 − 0.657

Notes: Only part of the results are selected and reported here; please refer to the cited papers for detail information. 2SLS: Two-stage Least Squares; HLM: Hierarchical Linear Modeling; CCR: Canonical Cointegrating Regressions.

The second issue that may influence results is the effect of internal migration[27]. The relatively healthy working-age labour from the relatively poor inland areas are more likely to self-select into working in the richer coastal regions. This may mean there is systematic measurement error in the health capital proxy. Some comfort is provided in the results of

[27] The following discussion also applies to education capital.

Hassler and Kuzin (2009), who show that in cointegration analysis, with procedures that account for short-run dynamics (e.g. DOLS) also correct for errors-in-variables, and are thus robust to measurement errors.

To summarise, the empirical results obtained in this chapter suggest that health investment needs to be a key component in the strategy for improved economic performance and overall development. The economic returns for such investment can be much greater for less developed regions and countries.

Chapter 4 will move on to investigate the short-run behaviour of health investment with regard to economic fluctuations from both public and private perspectives. Short-run health investment behaviour analysis is crucial as it can help in understanding the relationship between economic fluctuations and health status, which will be explored in detail in Chapter 5.

CHAPTER 4

THE CYCLICAL BEHAVIOUR OF PUBLIC AND PRIVATE HEALTH EXPENDITURE IN CHINA

4.1. INTRODUCTION

The well established two-way health-growth relationship suggests that in the long-run, the health status of a nation improves with economic growth, while at the same time, improved health status can make a positive contribution to economic growth (see Chapter 2 for more discussion). The above relationships are found to be non-linear, where the correlations are stronger in less developed, compared to more developed regions (see discussion in Chapter 3 for details). In empirical studies, the connections between health outcomes (or health inputs) and economic growth are estimated using different techniques, depending on the availability of data. For example, if a long-run time series dataset is available, a cointegration analysis technique is an appropriate method since macro time series are usually found to be nonstationary. On the other hand, when a (non-continuous) cross-country dataset is considered, standard panel data analysis techniques are usually adopted. In empirical studies, depending on data availability, four-year average or five-year average values are often used to avoid cyclical fluctuations (Rivera and Currais, 2004; Hartwig, 2010).

However, the relationship between economic fluctuations and health outcomes are complex (see Suhrcke and Stuckler (2012) for a detailed review). In developed countries economic recessions are usually associated with decreased mortality, while economic expansions are correlated with increased mortality (e.g. Tapia Granados, 2005). In developing countries, the reverse is often the case: economic recessions correspond to increased mortality (e.g. Cruces *et al*., 2012).

How health spending behaviour changes in response to economic fluctuations is one key channel linking economic fluctuations to health outcomes in developing countries (Waters *et al*., 2003; Hopkins, 2006; Gottret *et al*., 2009; Anderson *et al*., 2011). There are four important pathways that link government spending behavior changes to economic fluctuations in developing countries. First, falling government income through decreasing real tax revenue (adjusted by the inflation rate) may lead to cuts in government expenditure. Political distortion is the second possible factor: Talvi and Végh (2005) found that under optimal fiscal policy, political pressures could hamper running budget surpluses during economic booms. The inability to generate sufficiently large surpluses during expansions could in turn force the government to borrow less during recessions so as to satisfy the solvency constraint. The third reason that government could reduce spending during economic downturns is the existence of credit constraints during bad economic times - borrowing constraints can lead to governments being unable to implement countercyclical fiscal policies (Gavin and Perotti, 1997). The last factor is directly related to government health expenditure. During economic crises, foreign aid is likely to decline, although evidence of this connection is mixed (Stuckler *et al*., 2011). A substantial drop in foreign aid volume is likely to result in cuts to funding of health programs[28], many of which target poor people (USAID, 2009). Economic fluctuations could also impact on private spending behavior through an impact on purchasing power. The unemployment rate usually rises during economic recessions and job losses are likely to directly lower individual or household capacity to fund quality health care or adequate nutritional intake. In addition, increased prices of drugs and/or medical services[29] can lead to further pressure on private health financing.

Following Chapters 2 and 3, which investigate the long-run relationship between health investment and economic output in China, this chapter moves on to study the nexus between economic fluctuations and health spending behaviour from a short-run perspective. As discussed above, both long-run and short-run relationship studies are important components to understand the complex two-way health-growth relationship. Furthermore, these two groups of analyses will complement each other, providing additional information from a policy perspective.

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[28] As noted in Gottret *et al*. (2009, page 32), "in Rwanda and Ethiopia, over 50% of total government budgeted health expenditure is financed by donors, and off-budget donor funding for health accounts for more than 100% of government health expenditures."

[29] For example, there was a 67% increase in the price of medical treatment at Indonesian government health centres during the Asian economic crisis (Waters *et al.*, 2003).

Studies related to the cyclicality of government health expenditure can be traced back to a more general interest in total government expenditure. The cyclicality of fiscal policy has received increased attention by both researchers and policymakers since Gavin and Perotti (1997) found a different pattern of public spending in Latin America compared to more industrialised economies. Instead of increasing spending, public spending in Latin America declined during recessions, a sign of pro-cyclical fiscal behaviour. Numerous empirical analyses have since been conducted and their results suggest that fiscal policy tends to be counter-cyclical or acyclical in developed countries, but it is more likely to be pro-cyclical in developing countries (Lane, 2003). Furthermore, Gottret *et al*. (2009) warns that pro-cyclical public health spending behavior in developing countries might have a negative impact on the health-specific Millennium Development Goals (MDGs), especially for the poor (government expenditure on health is especially vulnerable during times of fiscal stress).

Although there are a number of empirical studies explaining the determinants of health expenditure (Newhouse, 1977; Gerdtham and Jönsson, 2000; to name but two), work on the cyclicality of government health expenditure is still a relatively new field. Darby and Melitz (2008), using data from 21 Organisation for Economic Co-operation and Development (OECD) countries, found that health-related public expenditures are counter-cyclical. On the other hand, Arze del Granado *et al*. (2010) found, using data from 150 countries, that public expenditure on health is acyclical in developed countries and pro-cyclical in developing countries. The limited empirical evidence suggests that the cyclical characteristics of health spending behaviour requires additional new research.

This chapter differs from the previous literature in two distinct ways. Firstly, rather than undertake further cross-country analysis this study focuses solely on China, the world's largest developing country. This has the benefit of reducing any bias due to heterogeneity and different approaches to data measurement across countries that may make unified models potentially implausible (Parkin *et al*. 1989; Kanavos and Mossialos, 1999; Melberg, 2011). Secondly, this paper also provides an exploratory analysis of the cyclical behaviour of private (i.e. out-of-pocket (OOP) personal) health expenditure. The evidence from micro level data in China suggests that during shocks to employment, urban residents face sharp declines in health benefits (Giles *et al*., 2006). It is still unclear, though, how an individual's health spending behaviour responds to economic fluctuations. Addressing these questions will provide useful information to policymakers, in both China and other transitional countries, concerning what kinds of fiscal policies may be useful to adopt in the event of economic fluctuations.

The remainder of the chapter is structured as follows. It begins in Section 4.2 with a brief introduction to China's health care finance and expenditure transition since the economic reform. The economic and econometric framework and data are presented in Section 4.3. Section 4.4 describes the stylised facts resulting from the time-series decomposition analysis, while Section 4.5 outlines the regression results from panel data analyses. Conclusions are presented in Section 4.6.

4.2. HEALTH CARE FINANCE AND EXPENDITURE IN CHINA

4.2.1. Health Care Finance Transition

When China embarked on economic reform from a centrally planned economy to a market-oriented economy in the late 1970s, like other transitional economies the government experienced a drastic reduction in its revenue, which in turn reduced its capacity to fund health care (Yip and Hsiao, 2008). The government converted its central subsidy for local governments into a block grant, imposing a hard budget constraint that effectively transferred responsibility for health services to local government (firstly decentralised to the provincial level, and in turn to county level governments) (Ho and Gostin, 2009; Schiere, 2010). Health care finance has depended largely on the economic performance of local government ever since (Young, 1990).

Two consequences of the fiscal decentralisation[30] process were a trend toward a distorted pattern of health service provision and an increase in the number of uninsured citizens (Blomqvist, 2007). Under hard budget constraints and local fee-generated revenues, China has evolved from a system of public salaries, public budgets, and state clinics to a predominantly fee-for-services system that is “public” in name only (Ho and Gostin, 2009). For example, as a result of the insufficient subsidy for state-owned hospitals they have been required to financially support themselves (Wang, 2009). It is therefore not surprising that

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[30] Decentralization could be clearly divided into political decentralization and fiscal decentralization - China has remained highly political centralized throughout but experienced significant changes (decentralization and then recentralization) in fiscal terms; the Chinese system is more centralized in government revenues than in expenditures (Guo, 2008).

this "financial autonomy" has forced health providers to rely on the sale of their services as a means of survival. The consequences of this fiscal decentralisation are also documented in a pattern of emphasising curative services rather than preventive care, and a fragmented three-tier health care delivery system[31] with little coordination of care (Young, 1990; Yip and Hsiao, 2008). The funding for preventative health services steadily declined, such that both public and private clinics were unwilling to provide services without payment (Schiere, 2010). Meanwhile a high proportion of people in both urban and rural areas were uninsured. According to the National Health Survey conducted by the Ministry of Health of China, 76.4% of the population had no form of health insurance in 1998, a situation that remained largely unchanged until 2003 when the figure dropped by 6% to 70.3% (Ministry of Health of China, 2009). With reduced government subsidies to health care, and with many people in both urban and rural areas being uninsured, individuals were responsible for the majority of healthcare financing. The evidence from survey data also confirms that the decreasing role of public health care financing leads to increasing cost burdens on the individual, especially for the more vulnerable in the population (e.g. the rural poor, migrant workers) (Parker and Thornton, 2007).

The outbreak of severe acute respiratory syndrome (SARS) in 2003 marked a critical wake-up call for China concerning the importance of public health and the role of the state in ensuring health care for its citizens (Schwartz and Evans 2007). Since then health care issues have gained significant prominence in public discussion and debate. The reminder of this section will discuss trends in health expenditure that show the health care financing transition. The data source is described in Section 4.3.2.

[31] In urban areas, from the lowest-tier to the highest, the three-tier health care delivery system is mainly composed of community health units, district hospitals and the city/provincial hospitals; in rural areas, village clinics, township hospitals and county/provincial hospitals (Chan *et al*., 2008).

4.2.2. Total Health Expenditure

Figure 4.1 shows the increasing trend and the shifting compositions of real per capita total health expenditure from 1978 to 2009. In 1978, the per capita total health expenditure was around 11.45 Yuan (7.26 US$[32]) with government, social and private health expenditure accounting for 32.16%, 47.41% and 20.43% respectively. Real total health expenditure has shown an increasing trend ever since, but the roles of these three components change, with a decreasing proportion of government and social health expenditure, and an increasing role in OOP health expenses until 2001 and 2002. In 2002 the three components of total health expenditure, government, social and private, accounted for 15.69%, 26.59% and 57.72% respectively. In 2009 real per capita total health expenditure had grown to 248.36 Yuan (constant 1978 prices), and along with the rethink of the government role in health care financing after the SARS attack, and the re-introduction of basic medical insurance in both rural and urban China, the proportions of the three components, government, social and private, were 27.23%, 34.57% and 38.19% respectively.

As mentioned above, total health expenditure in China is categorised into three components by the Ministry of Health of China: government health expenditure, social health expenditure, and private health expenditure (i.e. OOP health expenses); see Zhao *et al.* (2004), Ministry of Health of China (2010), Pan and Liu (2012) for more detail. These three components are now examined in more detail.

[32] Calculated based on the average exchange rate in 1978: 1 US Dollar = 1.577 Chinese Yuan. The average exchange rate in 2009 is: 1 US Dollar = 6.831 Chinese Yuan.

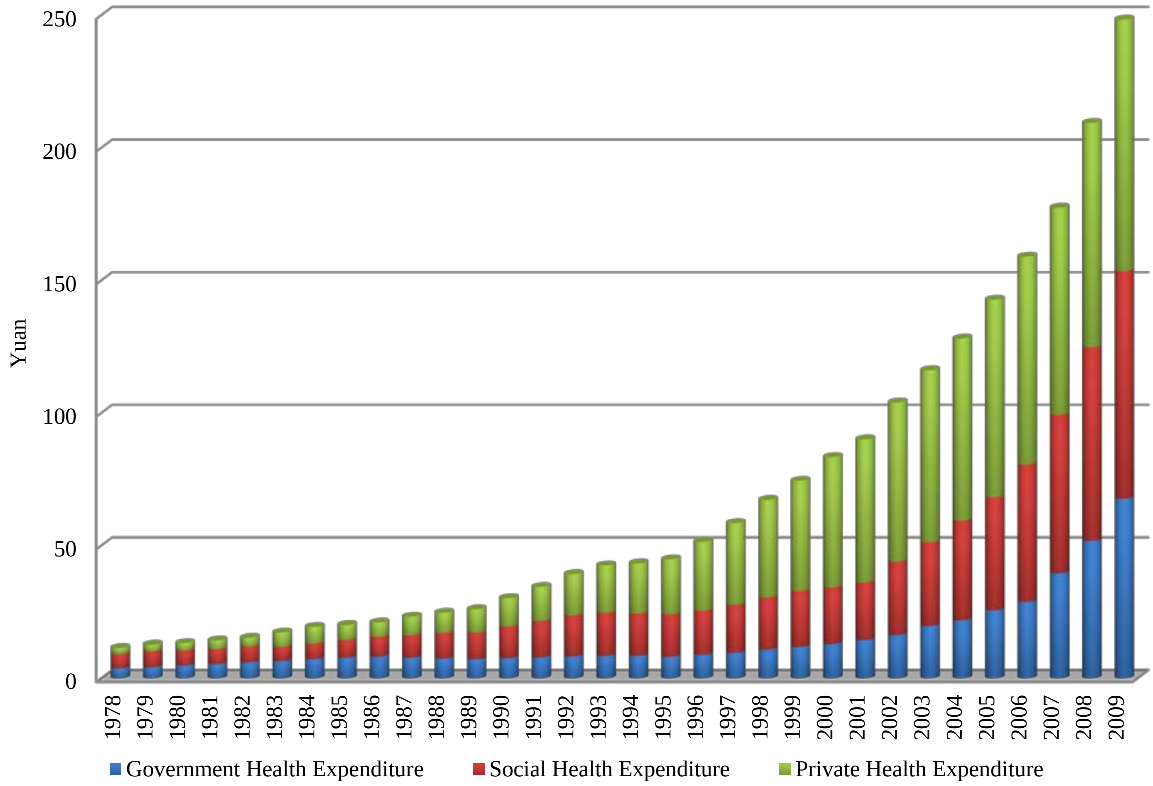

Figure 4.1. Real health expenditure per capita of China, 1978-2009 (Yuan, 1978 price).

4.2.3. Government Health Expenditure

The Ministry of Health of China (2010) groups government health expenditure into four components: health services expenditure, medical insurance fund subsidies, family planning expenditure and health administration expenditure. As can be found in Figure 4.2, health services expenditure is the biggest component, followed by medical insurance fund subsidies. Medical insurance fund subsidies used to be referred to as the Government Insurance Scheme (GIS) by government agencies; from 2003 onwards, this catalogue also includes the basic medical insurance fund subsidies[33].

[33] Calculations are based on data reported by the Ministry of Health of China (2009, 2010). In 2005 the basic medical insurance fund subsidies account for around 8% of medical insurance fund subsidies, 2% of government health expenditure.

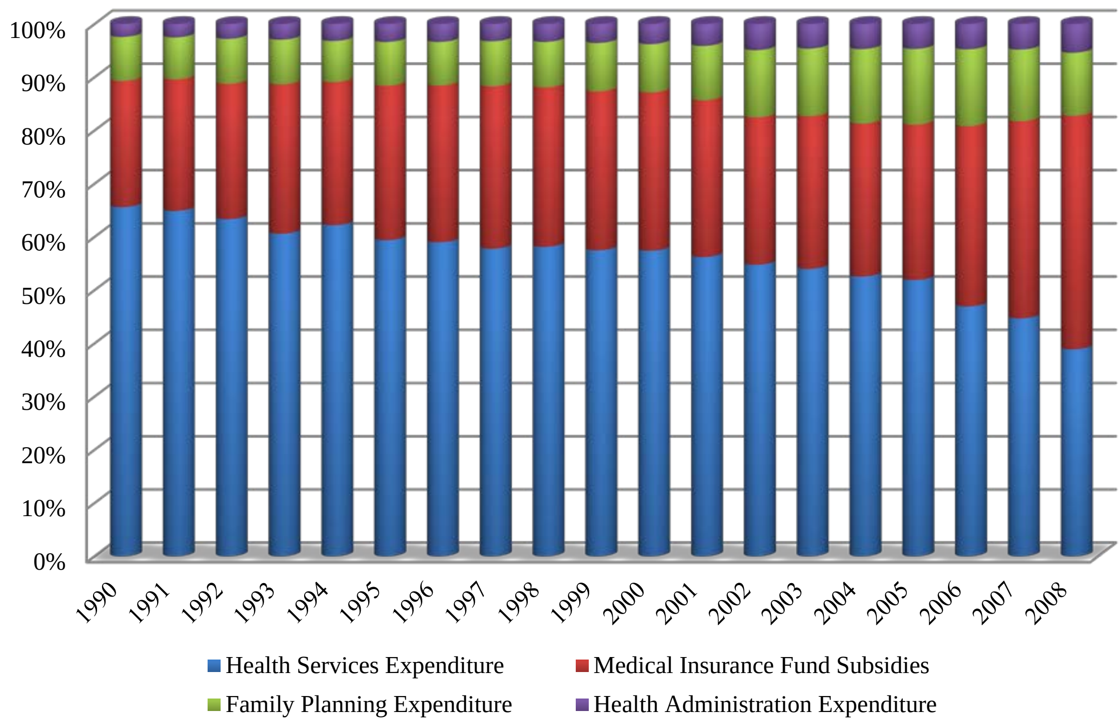

Figure 4.2. The composition of government health expenditure of China.

It is also worth noting that a distorted spending pattern can be observed in the monthly spending data. A plot of monthly government health spending for the years 2007 to 2010 reveals that on average, governments spend one quarter of their total budget in December (See Figure 4.3). This pattern is not specific to health spending: Dunaway and Fedelino (2006) found that a similar pattern can be observed for total government expenditure. This may cast doubt on whether government health expenditure has been used efficiently.

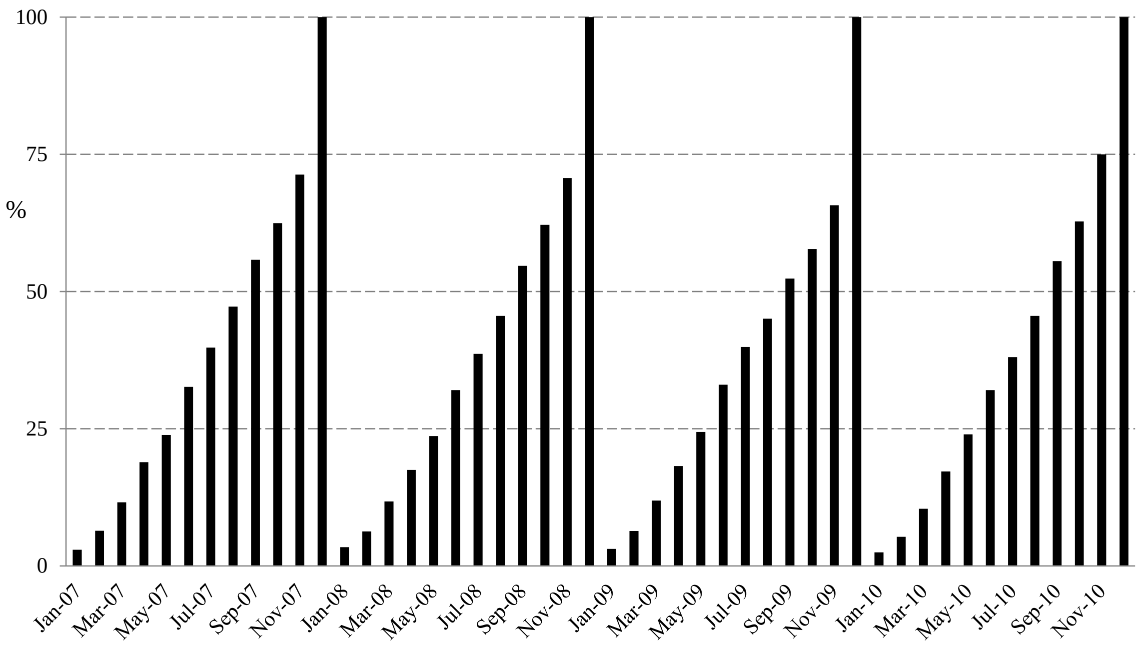

Figure 4.3. Accumulated monthly government health expenditure of China.

4.2.4. Social Health Expenditure

The second component of total health expenditure is social health expenditure; this mainly includes: (1) medical insurance premiums and non-government subsidies to the Labor Insurance Scheme (LIS)[34], the Urban Employee Basic Medical Insurance (UEBMI), the New Rural Cooperative Medical Scheme (NRCMS), and the Urban Resident Basic Medical Insurance (URBMI), (2) the extra-budgetary capital investment, and (3) private direct investment in practices (Zhao *et al*. 2004, page 6; Chan *et al*., 2008, page 132). The detailed composition of social health expenditure is not reported by the Ministry of Health of China. According to Zhao *et al*. (2004), in 2002, the medical insurance premiums and

[34] The LIS is the predecessor of the Urban Employee Basic Medical Insurance. It is financed by each enterprise, and under this medical insurance scheme employees and their dependents and retirees are all covered.

subsidies accounted for around 67.3% of total social health expenditure. As mentioned before, in 1978 social health expenditure accounted for 47.41% of total health expenditure. This figure reached its lowest in 2001 at less than one quarter of expenditure (24.10%). With the re-building of the medical insurance system from the beginning of the new century, the percentage of social health expenditure increased again and reached 34.57% in 2009 (although this is still less than in 1978).

The new wave of health care financing reform is characterised by the insurance-based model[35]. The government's aim is to expand public health insurance coverage with a target of achieving 90% health insurance coverage by the end of 2010, and universal coverage of primary health services by 2020 (Guo *et al*., 2010). According to the latest National Health Survey in 2008, the percentage of the population without any social health insurance has decreased sharply to 12.9% (Ministry of Health of China, 2009). One of the greatest merits to having universal coverage is that health insurance may reduce the monetary cost of health services, which in turn may increase access to health services for poor and underprivileged households[36] (Sepehri *et al*., 2006). However the empirical evidence regarding social health insurance in China seems to tell a different story. Wagstaff *et al*. (2009a), using national health survey data, found that in rural China enrolling in NRCMS significantly increased health services utilisation, however, with concurrent increases in their OOP expense per visit. The primary reason for this limited financial protection effect of social health insurance may be due to the supplier-induced demand that, owing to limited government subsidies, selling drugs is still a main source of income for hospitals. Other reasons may relate to the benefit

[35] The URBMI for urban residents, together with the UEBMI for urban employees and the NRCMS designed for rural populations, constitute a natural platform from which to merge these three existing schemes—leading to universal health insurance coverage for the people of China (Lin *et al*., 2009).

[36] The real situation is always more complicated than theory predicts. For example, Wagstaff (2010) shows that in Vietnam, although the OOP spending has been reduced significantly, the health services utilisation did not increase significantly. This result may indicate that insured people face multiple non-price constraints.

packages designed and the cautious attitude for spending the insurance funds by the local governments[37]. Based on the only available figures from urban basic medical insurance, the distribution of the percentage of the accumulated medical insurance balance as a share of current insurance fund expenditure, from 2001 to 2009, see Figure 4.4. It can be seen that both the mean and median value are higher than one hundred percent for all nine years. It is important to evaluate whether the introduction of medical insurance can effectively relieve the financial burden on the Chinese population.

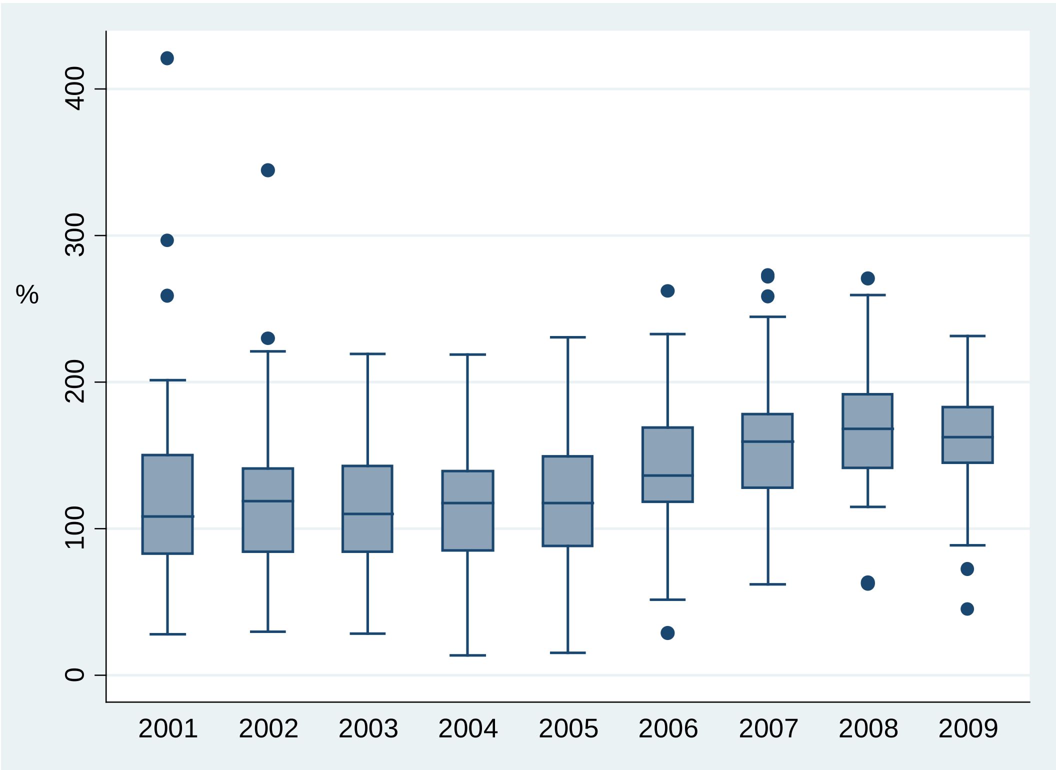

Figure 4.4. Urban basic medical insurance accumulated balance share of current expenditure.

[37] A case study by Zhang *et al*. (2010) investigated the balance of NRCMS funds and the influencing factors using data from six counties in two provinces. They found that although under the general expectation a surplus of 5-10% of total fund revenue should be maintained for unexpected risks, five out of six counties held a larger surplus (i.e. 28%-66% for the total fund).

4.2.5. Private Health Expenditure

The last component of total health expenditure is the private OOP health expenditure. As part of China's economic transition and health system reform, individuals have increasing responsibility for the majority of healthcare financing. As can be seen in Figure 4.1, real private health expenditure keeps increasing and its proportion peaked at around 60% of total health expenditure in 2001. On further comparison of per capita private health expenditure in rural versus urban areas, two main patterns can be observed in Figure 4.5. Firstly, private health expenditure as a percentage of total personal expenditure has an upward trend for both urban and rural residents. It increased from 2.48% in 1985 to 6.98% in 2009, with a peak at 7.57% in 2005 for urban residents. For rural residents, although there are some fluctuations, the figure increased to 7.20% in 2009, from a low 3.93% in 1985. Secondly, per capita private health expenditure increased much faster for urban residents than rural residents until around 2002. An empirical analysis of the relationship between basic medical insurance expenditure and the private OOP health expenditure would tell us whether the expansion of medical insurance coverage can help relieve the financial burden on the Chinese population.

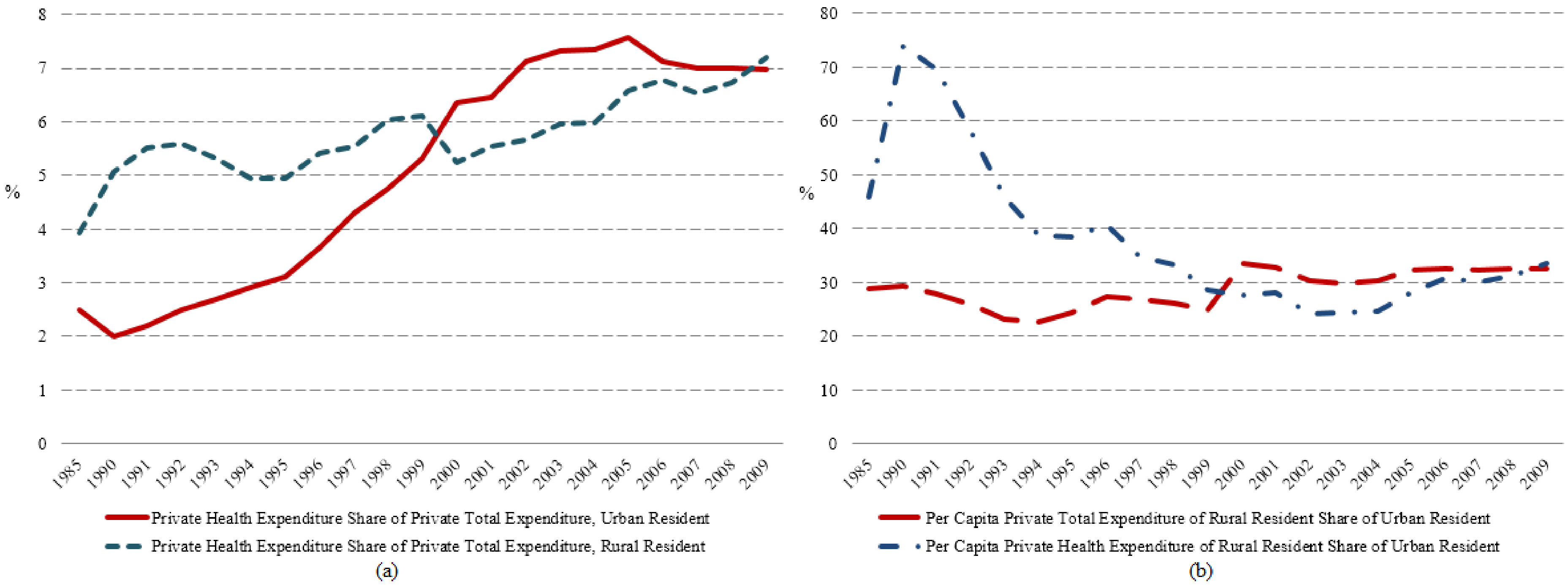

Figure 4.5. Comparison of urban-rural private health expenditure.

4.3. METHODOLOGY AND DATA

4.3.1. Economic and Econometric Framework

In this chapter, two groups of data (national time series data and provincial-level panel data) are used to analyse the cyclical behaviour of health expenditure. Accordingly, two different econometric methods are adopted for the empirical analyses. Firstly, national time series data are decomposed into permanent and cyclical components using a multivariate Beveridge-Nelson technique. The stylized facts are then presented using correlation analysis on the decomposed cyclical components. Secondly, panel data models are used to study the cyclical behavior at provincial level, with instrumental variables (IV) models allowing causal relationships to be identified in some cases.

4.3.1.1 National time series analysis

There are several methods for decomposing time series into permanent and cyclical components, see Mills (2003) for a review. The permanent line is the domain of growth theory and thought to be determined by long-run factors while the cyclical component is related to business cycles, which both monetary and fiscal policies try to eliminate (Enders, 2010). The cyclical element is of interest in this study. In this chapter, a multivariate Beveridge-Nelson (BN) approach (a generalization of the method outlined in Beveridge and Nelson (1981)) proposed by Garratt *et al*. (2006a) and Garratt *et al*. (2006b) is preferred assuming a cointegrated relationship can be justified. In contrast to the classical univariate filter (e.g. the Hodrick-Prescott (1997) filter), the key benefits of this multivariate decomposition strategy is that it allows for long-run restrictions, as well as short-run interactions that might exist among a group of cointegrated series. An introduction to this multivariate BN decomposition approach is available in Appendix 4.

As stated above, the existence of a cointegration relationships among the investigated series is necessary to adopt this multivariate BN decomposition approach[38]. Analysis starts by examining whether the data are stationary or not. In addition to the Phillips and Perron (1988) unit root test, which does not consider potential structural breaks, the Perron and Vogelsang (PV) (1992) single structural break unit root test has also been adopted. Specifically, the PV test allows for either a gradual shift in the mean of the variable (innovational outlier, IO case) or a sudden change in a series (additional outlier, AO case). If a unit root exists, a Vector Error Correction (VEC) modeling framework is further considered. To determine the number of long-run cointegration relationships among the non-stationary series, the Johansen cointegration test (the multivariate test based on the autoregressive representation explored by Johansen (1988)) is used. Next, a VEC model is estimated, in which the short-run behavior of the investigated series is tied to long-run values through an error-correction process. To avoid the possibility of quadratic trends in some of the series, the trend is restricted and the intercept is unrestricted. Finally, the decomposed permanent and cyclical series are attained according to the method documented at Garratt *et al*. (2006a) and Garratt *et al*. (2006b) (see Appendix 4.1 for a brief summary). The empirical analysis is performed by Microfit Version 5.0 software.

4.3.1.2 Provincial-level panel data analysis for government health expenditure

The above multivariate decomposition analysis provides a broad picture of the relationships

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[38] If there is no cointegration relationship, then the results will be identical to the one based on univariate BN decomposition technique (Garratt *et al*, 2006b).

among cyclical components in this series; however, no causality can be drawn from the above correlation analyses. This results may be due to omitted variable bias or reverse causality bias. The second part of the econometric analysis, which is based on provincial-level panel data, addresses this issue in more detail.

In a standard macro health expenditure equation (e.g. Pan and Liu, 2012), government health expenditure is driven mainly by three groups of factors: national income, demand/need factors, and other social factors. The equation of interest is:

$$\ln(GHE_{it}) = \alpha_0 + \alpha_{1i} t + \beta_1 \ln(GRP_{it}) + \gamma_1' x_{1it} + u_{0it} \quad (4.1)$$

$$u_{0it} = \mu_{0i} + \varepsilon_{0it} \quad (4.2)$$

where ln(GHE) and ln(GRP) are the natural logarithm of real government health expenditure and real gross regional product (GRP) respectively, t is a province specific linear time trend, x_1 is a vector that contains other relevant control variables (e.g. demand/need factors and other social factors), based on previous literature and the availability of data in China, the subscript i and t denote the province and time period respectively, μ_{0i} captures time-invariant unobservable heterogeneity, ε_{0it} is an idiosyncratic error term. It is assumed that μ_{0i} is uncorrelated with ε_{0it}, and may be correlated with the other regressors in Equation (4.1). Pan and Liu (2012) adopt a similar framework to study the level of government health expenditure in China.

Since the main aim of this study is to examine the cyclical behaviour of health expenditure, the first difference of Equation (4.1) is taken:

$$\Delta \ln(GHE_{it}) = \alpha_{1i} + \beta_1 \Delta \ln(GRP_{it}) + \gamma_1' \Delta x_{1it} + \varepsilon_{1it} \quad (4.3)$$

where Δ is the first difference operator, α_{1i} is the time-invariant unobservable heterogeneity for each province, $\varepsilon_{1it} = \varepsilon_{0it} - \varepsilon_{0it-1}$. The province-specific α_{1i} can be eliminated by using fixed effects (FE) estimation. For the idiosyncratic error term ε_{1it}, two key assumptions are tested: whether the variances of the errors across provinces are identical (i.e. homoskedasticity) and whether the error term is serially uncorrelated. The first assumption may not hold since in this analysis, the cross-sectional units (provinces) have very different populations. The assumption is tested using a modified Wald statistic for groupwise heteroskedasticity in the residuals of a fixed effect regression model, proposed by Greene (2002) and coded as "xttest3" in Stata. The second assumption is tested using a Wald test proposed by Wooldridge (2002) and coded as "xtserial" in Stata (Drukker, 2003). The heteroskedasticity and autocorrelation tests results suggest that the null hypothesis of homoskedasticity is rejected at 1% level and the null of no serial correlation cannot be rejected at 10% level. Based on above discussion, the heteroskedastic robust standard errors are used.

The coefficient β_1 provides a measure of cyclicality of government health expenditure: it measures the elasticity of expenditure with respect to output growth. A significant positive value of β_1 implies pro-cyclical behaviour (a value above unity implies a more-than-proportionate response to output fluctuations); while a significant negative value implies counter-cyclical behaviour. The analysis framework is similar to empirical studies which use a cross-country data set to investigate the cyclicality of government health expenditure or government expenditure (e.g. Arze del Granado *et al*., 2010; Erbil, 2011).

An additional concern here is whether to include the lagged dependent variable in Equation (4.3). Erbil (2011) suggests that the lagged dependent variable should be included to allow for the long-term mean reversion in fiscal behaviour, since the last years' policy decisions or the shock to the fiscal balance may have a lag effect. Following this logic, a second empirical setting for the cyclical behaviour of government health expenditure equation can be specified as:

$$\Delta \ln(GHE_{it}) = \alpha_{1i} + \sum_{l=1}^{\max lag} \rho_l \Delta \ln(GHE_{i,t-l}) + \beta_1 \Delta \ln(GRP_{it}) + \gamma_1' \Delta x_{1it} + \varepsilon_{1it} \quad (4.4)$$

A range of control variables are included in the x_1 vector. First, the natural logarithm of real central government transfer (ln(TRANSFER)) is included to capture effects of central government redistribution of tax revenue. According to the Ministry of Finance of China (2009), in 2008, on average 38% of total local government expenditure comes from central government transfers; in inland China, more than half (54.4%) of local government expenditure relies on central government transfers. The detailed central government transfer components are unavailable for the sample years. The latest figure in 2011[39] suggests that the local government's health expenditure was 629.62 billion Yuan, in which 167.65 billion Yuan (26.6%) comes from central government transfers. According to the Minister of Health of China, Mr. Chen Zhu, the majority of the central government's transfer expenditure on health focuses on inland China[40].

Secondly, four health status related variables are included, to capture the demand for health care: the crude death rate (CDR), the percentage of population aged between 0 and 14 (POP014), the percentage of population aged 65 and older (POP65P), and a time dummy

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[39] http://www.npc.gov.cn/pc/11_5/2012-03/19/content_1715336.htm.

[40] http://news.xinhuanet.com/politics/2007-09/05/content_6666357.htm.

for the year 2003 (D2003) (which is included to capture the national SARS shock)[41]. Thirdly, the urbanisation rate (URB) is included to capture a price effect that usually occurs for health services, whereby the price in urban areas is higher than rural areas. As a robustness check, models are estimated with these covariates included in level forms instead of the preferred first difference form.[42]

A natural starting point to estimate Equation (4.3) is via a FE estimator. However, if the model is to be used to provide causal explanations, the issue of endogeneity needs to be addressed. The first source of endogeneity relates to reverse causality: human capital theory suggests that expenditure on health will influence a population's health status and further impact on labour productivity or aggregate output (see Appendix 2 for a review). For example, using data from OECD countries, Beraldo *et al*. (2009) found that both public and private health expenditure have a significant positive effect on GDP growth. This reverse causality would mean the FE estimates would overstate the causal effect of growth on health expenditure. Reverse causality could also occur because of more general fiscal policy effects. While government expenditure is influenced by contemporaneous economic output, at the same time the expansionary effect of government spending will also impact on economic output (e.g. Ilzetzki and Vegh, 2008). This provides another source of upward bias in estimating equation (4.3) via FE. However, there is some evidence from developing country data showing that government health spending has had

[41] The unit root test in the time series analysis also identified a significant structural break for public health expenditure growth rate at 2003. See Table 4.4.

[42] Aside from the theoretical reason to include those variables in first differenced term (as shown in Equation (4.3)), an additional benefit of using this framework compared to the levels form is that usually the first difference of a macro-time series is stationary (since macro time series is usually integrated of order one). Since the time series in the panel data analysis is relative short, it would be difficult to draw a solid conclusion using a unit root test.

negative effects on GDP growth rates (Devarajan *et al*., 1996; Ghosh and Gregoriou, 2008). Devarajan *et al*. (1996) study the composition of central government public expenditure and economic growth in 43 developing countries from 1970 to 1990. They found the effect of public health spending is negative but insignificant, relative to economic growth. Ghosh and Gregoriou (2008) further study the link between government expenditure components and economic growth in the context of optimal fiscal policy. By using data from 15 developing countries from 1972 to 1999, Ghosh and Gregoriou confirm the results reported in Devarajan *et al*. (1996) that the coefficient of health component is significantly negative. If this was the case, there would be a downward bias in the FE estimator.

For China, it is unclear which of the above scenarios might dominate. On one hand, the positive effect of health investment on economic growth via human capital theory may only become apparent in the long-run. On the other hand, an empirical study investigating the fiscal policy effect by Zhang and Zou (1998) using province level data in China (1987 to 1993) found no significant association between government spending on human capital (measured as the share of budgetary spending on culture, education, health, and science out of total budgetary spending) and economic growth. To summarise, although it is possible that endogeneity is a potential issue in this study, it is unclear whether ignoring this issue will under- or over- estimate the true causal effect. Results are reported in the paper using instrumental variables, the Hausman test for endogeneity will provide some evidence about the presence or magnitude of endogeneity bias.

The IV approach, using a generalised method of moments (GMM) estimator, is adopted, using the one-period lagged value of potentially endogenous variables as an instrument for current value, consistent with prior literature. The Δln(GRP) (hereafter GRP growth or economic growth) in the current year is instrumented by one year lagged Δln(GRP). Similarly, the change in the crude death rate (hereafter mortality growth ΔCDR) could also be endogenous due to possible reverse causality, and so it is instrumented by the one year lagged ΔCDR. While the lagged values may not be ideal instruments, this is the most commonly adopted approach in applied macro analyses. In cross-country analysis, lagged GDP growth is often used as an instrument for current GDP growth (e.g. Lane, 2003; Galí and Perotti, 2003; Arze del Granado *et al.*, 2010). Another commonly adopted IV for GDP growth is the weighted GDP growth of each country's trading partners. However this approach is difficult to adopt in this provincial-level analysis.

One additional instrument for GRP growth, as used by He (2011) and inspired by Miguel *et al*. (2004), has also been considered: the yearly temperature difference (TempDiff) of each capital city within the 31 provinces. The role of domestic agricultural production in the maintenance of an adequate growth rate for China is crucial: evidence suggests that China has experienced many economic cycles and all the poor years were related to inadequate agricultural performance (Yao, 2000). Weather shocks are one key factor that exogenously impact on the agricultural sector in both developing and developed countries (Miguel *et al*., 2004; Schlenker and Yields, 2006; Deschênes and Greenstone, 2007; Barrios *et al*., 2010; Dell *et al*., 2012). Studying household production in Shandong Province (China's largest agricultural province), Zhang (2008) noted that severe weather shocks can increase poverty rates; furthermore, among the various risk factors associated with agriculture (e.g. prices change, quality of agriculture input), weather shocks have the most extensive impacts. The impact of weather shocks on the agricultural sector in China has also been highlighted in Boyd *et al*. (2011) and Xu *et al*. (2011). Based on the above discussion, it is reasonable to argue that weather shocks provide a relevant instrument because of their negative impact on the agricultural industry and thus on GDP growth in China. There could still be some concern about whether weather shocks can be legitimately excluded from the expenditure equation: evidence suggests weather shocks may also have a negative impact on health status and thus increase health expenditure. However, considering the changing health status is already partly controlled for in this study (through the change in crude death rates), this exclusion restriction should be valid. The yearly temperature difference (TempDiff) used to measure weather shocks is calculated as the difference between the highest and lowest monthly average temperatures (He, 2011).

The second problem which makes an OLS estimator inconsistent is related to estimating Equation (4.4) when a lagged dependent variable has been included. Given the above issues and following the literature (e.g. Baltagi, 2005; Cameron and Trivedi, 2010), a two-step GMM estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998) is adopted for the dynamic panels.

4.3.1.3 Provincial-level panel data analysis for private health expenditure

The empirical research studying private (OOP) health expenditure mostly uses survey data. A key research question in this context is to see if the introduction of medical insurance can have a financial protection effect on private health expenditure. In this chapter aggregate data are used to study the cyclicality of private health expenditure, and also the potential medical insurance effect. Given that provincial level basic medical insurance expenditure data is only available for urban residents, this component on the study will focus on urban samples only.

The main equation for private health expenditure is specified as:

$$\Delta \ln(PHE_{it}) = \alpha_{2i} + \beta_2 \Delta \ln(INCO_{it}) + \beta_3 \Delta \ln(BMI_{it}) + \gamma_2' \Delta x_{2it} + \varepsilon_{2it} \tag{4.5}$$

where *PHE* is per capita real private health expenditure of urban residents, *INCO* is per capita real disposable income[43] of urban residents, *BMI* is real urban basic medical insurance (i.e. UEBMI and URBMI) expenditure, α_{2i} is the time-invariant unobservable heterogeneity for each province, ε_{2it} is an idiosyncratic error term. As with the GHE equation, the other control variables include the crude death rate, the percentage of population aged between 0 and 14, the percentage of population aged 65 and older, a time dummy for the year 2003 and the urbanisation rate. Government health expenditure is also considered to see if private health expenditure may respond to changes in the level of government health expenditure. Central government transfers is not included here since

[43] According to the National Bureau of Statistics, the formula for calculating disposable income is "disposable income = total household income - income tax - personal contribution to social security - subsidy for keeping diaries for a sampled household".

this should not have a direct effect on private spending behaviour; its indirect effect will be captured by the newly added government health expenditure variable.

Finally, the role of fiscal policies may be different between China's two broad regions (inland provinces and coastal provinces) since the central government aims to use central transfers to reduce the disparities in fiscal capacity across regions (Huang and Luo, 2009). Following the above considerations all analyses are therefore undertaken on the whole sample, and two sub-samples (inland and coastal China samples).

4.3.2. Data

The national time series data ranges from 1978 to 2009; the provincial-level panel data for GHE equation starts from 1996, and the PHE equation only uses data from 2001 due to limited availability of urban basic medical insurance expenditure data.

One key issue that needs to be emphasised here is the measurement of health expenditure. As mentioned earlier, national total health expenditure is reported in three categories by the Ministry of Health of China in the *China's Health Statistical Yearbook*: government health expenditure, social health expenditure and private health expenditure. In the time series analysis, following Jeong (2005), a new variable called public health expenditure is generated by combining government health expenditure and social health expenditure approximating according to OECD standards.

In the panel data analysis, the provincial government health expenditure data is drawn from the *Finance Yearbook of China* across several years, compiled by the Ministry of Finance of

China. The provincial data for per capita private health expenditure of urban residents is drawn from the *China Statistical Yearbook* compiled by the National Bureau of Statistics of China. The provincial basic medical insurance expenditure data for urban residents are available from the *China Labour Statistical Yearbook* compiled by the National Bureau of Statistics of China, and the Ministry of Labour and Social Security of China since 2001.

4.3.2.1. National time series data

Real GDP is calculated by combining the nominal value of GDP in 1978, and the GDP index. Real health expenditure data are adjusted by the General Consumer Price Index (CPI). As is common in the time series literature, per capita forms of these variables are constructed using the total population at year end. All variables are transformed into natural logs. The summary statistics for the national time series and also the decomposed permanent and cyclical components are presented in Table 4.1.

Table 4.1. Summary statistics for national time series and the decomposed components.

Variable	Mean	Std. Dev.	Min	Max	Obs.
ln(GDP per capita)	7.173	0.789	5.937	8.534	32
permanent component	7.253	0.743	6.036	8.515	30
cyclical component	-0.001	0.032	-0.080	0.041	30
ln(Public health expenditure per capita)	3.277	0.760	2.209	5.034	32
permanent component	3.344	0.760	2.390	5.174	30
cyclical component	0.001	0.123	-0.220	0.306	30
ln(Private health expenditure per capita)	2.856	1.205	0.850	4.552	32
permanent component	2.987	1.127	1.059	4.467	30
cyclical component	0.000	0.095	-0.191	0.208	30

Note: The permanent and cyclical components are calculated through multivariate BN decomposition approach.

4.3.2.2. Provincial-level panel data

Real GRP is calculated by combining the nominal value of GRP in 1978, and the GRP index. Real health expenditure data and real central government transfer data are adjusted by the provincial CPI. Unless specified, all data are drawn from the *China Statistical Yearbook* over several years. The summary statistics for the provincial-level panel data are presented in Table 4.2.

Table 4.2. Summary statistics for provincial-level data†.

Variable		Mean	Std. Dev.	Min	Max	Obs.	Year
Δln(Real government health expenditure, GHE)	overall	0.177	0.140	-0.228	0.718	401(a)	1997-2009
	between		0.021	0.135	0.212	31	
	within		0.138	-0.198	0.748	13	
Δln(Real gross regional product, GRP)	overall	0.108	0.022	0.046	0.213	434	1996-2009
	between		0.010	0.089	0.142	31	
	within		0.019	0.051	0.180	14	
Δln(Real central government transfer income, TANSFER)	overall	0.171	0.125	-0.364	0.578	401(a)	1997-2009
	between		0.043	0.056	0.246	31	
	within		0.118	-0.392	0.539	13	
Δ(Crude death rate, CDR)	overall	-0.045	0.332	-1.560	1.120	432(b)	1996-2009
	between		0.057	-0.266	0.031	31	
	within		0.327	-1.532	1.171	14	
POP014	overall	0.208	0.052	0.076	0.351	403	1997-2009
	between		0.044	0.105	0.274	31	
	within		0.029	0.136	0.289	13	
POP65P	overall	0.082	0.020	0.041	0.164	403	1997-2009
	between		0.017	0.054	0.137	31	
	within		0.011	0.058	0.109	13	
Urbanization rate	overall	0.406	0.151	0.166	0.883	403	1997-2009
	between		0.145	0.201	0.804	31	
	within		0.051	0.254	0.531	13	
TempDiff	overall	25.884	7.539	9.200	47.300	403	1997-2009
	between		7.433	11.377	40.731	31	
	within		1.800	19.953	32.453	13	
Δln(Per capita real private health expenditure of urban residents, PHE)	overall	0.091	0.117	-0.442	0.577	248	2002-2009
	between		0.028	0.038	0.144	31	
	within		0.113	-0.389	0.526	8	
Δln(Per capita real basic medical insurance expenditure of urban residents, BMI)	overall	0.315	0.317	-0.296	3.052	247(c)	2002-2009
	between		0.108	0.122	0.570	31	
	within		0.299	-0.452	2.796	8	
Δln(Per capita real disposable income of urban residents, INCO)	overall	0.087	0.031	-0.072	0.190	248	2002-2009
	between		0.015	0.047	0.112	31	
	within		0.028	-0.031	0.231	8	

Note: Unbalanced panel ((a) The data of the Municipality of Chongqing and Sichuan Province in 1997 are missing; (b) the data of Shaanxi Province in 2000 and 2001 are missing; (c) the data of Tibet Autonomous Region in 2002 is missing). † Between: variation in mean across provinces. Within: deviations of each province from its own average.

4.4. STYLISED FACTS

In this section, the stylized facts are presented through correlation analysis on the decomposed cyclical health expenditure series. A cointegrated relationship is a prior requirement for using the multivariate BN approach. The empirical analysis begins by identifying whether those series are cointegrated. Starting from the unit root tests, as can be found in Table 4.3, without considering the structural breaks, except public health expenditure which seems to be integrated of order two ($I(2)$), all other series are integrated of order one.

Table 4.3. Phillips-Perron unit-root test results.

Variable	Z(rho)	Z(t)	Trend
Level form			
ln(GDP per capita)	-10.683	-2.345	YES
ln(Public health expenditure per capita)	4.381	2.436	YES
ln(Private health expenditure per capita)	-6.425	-1.489	YES
First difference form			
Δ ln(GDP per capita)	-12.911**	-2.859**	NO
Δ ln(Public health expenditure per capita)	-8.762	-1.984	NO
Δ ln(Private health expenditure per capita)	-27.386**	-5.514**	NO

Notes: ** indicates the significance level at 5%. The Newey-West lags is set based on int{4(T/100)^(2/9)}. Whether a linear trend is included (YES: included; NO: not included) for each variable is decided by economic theory and line graph.

However, there is likely to be a structural break in the public health expenditure series because of the outbreak of SARS. The policy response to SARS may have induced a sudden change in the series and thus an AO case would be more relevant[44]. That noted, the

[44] Assuming the IO case will not change the conclusion for the order of integration for each series. However,

unknown structural break for real public health expenditure growth identified by the PV unit root test is in year 2003 (see Table 4.4 for details). This information is used in the panel data analysis via the inclusion of a time dummy for the year 2003 to capture the health crisis (SARS) shock. Given the potential structural breaks of each series, it can be concluded that all series are $I(1)$.

Table 4.4. Perron-Vogelsang one break unit root test results (additional outlier case).

Variable	Lags	TB_1	DU_1	$(\rho$-1)
Level form				
ln(GDP per capita)	0	1997	1.342**	-0.195
ln(Public health expenditure per capita)	0	2004	1.585**	-0.202
ln(Private health expenditure per capita)	0	1993	2.089**	-0.297
First difference form				
Δ ln(GDP per capita)	1	1987	0.005	-0.619**
Δ ln(Public health expenditure per capita)	0	2003	0.114**	-0.774**
Δ ln(Private health expenditure per capita)	3	1981	-0.004	-0.461

Notes: Δ is the first difference operator. TB_1 is the time period when the mean is being modified; DU_1 is a sustained dummy variable capturing a mean shift occurring at the break date; ρ is autoregressive parameter. ** indicates the significance level at 5%. Lags included were selected following the procedure suggested in Clemente *et al*. (1998) with a max lag of 5.

In the next step, the Johansen cointegration test has been adopted to test whether the above series are cointegrated or not. Considering the time period and information criteria (e.g. Schwarz Bayesian Criterion), the lag for the VAR is set to be 2. The cointegration LR test results are mixed: (1) the trace of the stochastic matrix suggests that there are two cointegration (rank = 2) relationships at 5% criterion; (2) the maximal eigenvalue of the stochastic matrix suggests that there are no cointegration relationships (rank = 0) at the 10% criterion. On the other hand, the choice of the number of cointegrating relationships using model selection criteria (i.e. Akaike Information Criterion, Schwarz Bayesian Criterion,

the structural breaks found based on AO case are more close to the documented history events and policy change.

and Hannan-Quinn Criterion) suggests that rank = 3. Considering the well documented cointegration relationship between GDP and total (or subcomponents of) health expenditure (e.g. Gerdtham and Löthgren, 2000; Clemente *et al*. 2004; Sharma and Srivastava, 2011), there should be at least one cointegration relationship. This hypothesis is tested by the Engle-Granger residual based approach. The residual from the first step equation (irrespective of whether a linear time trend is included or not) is stationary, suggesting a cointegration relationship does exist. Combining all the above results, it is concluded that there is at least one cointegration relationship among the three series.

The multivariate BN decomposition analysis is conducted through the VEC model. The cointegrated series are decomposed into permanent (trend) and cyclical (transitory) components.

For the permanent components, it can be seen clearly from Figure 4.6 that when excluding the influence of business cycles, private health expenditure exceeds public health expenditure after 1993 and this trend persists until 2005. The key factor influencing the first turning point may be related to the central government's attempt to introduce market forces into the health system. Followed by a series of State Council Policy Documents, for example *Report on Issues Concerning Health Reform* (State Council Policy Document 1985 No. 62), *Opinions on Expanding Health Services* (State Council Policy Document 1989 No. 10), and *Opinions on Deepening the Reform of the Health System* (Ministry of Health Policy Document) in 1992, the public health institutions were gradually recognised as economic bodies instead of pure welfare entities (Liu and Mills, 2002). The ideology of these documents suggests that in China, financing of the health care system is shifting from central government to the individual.

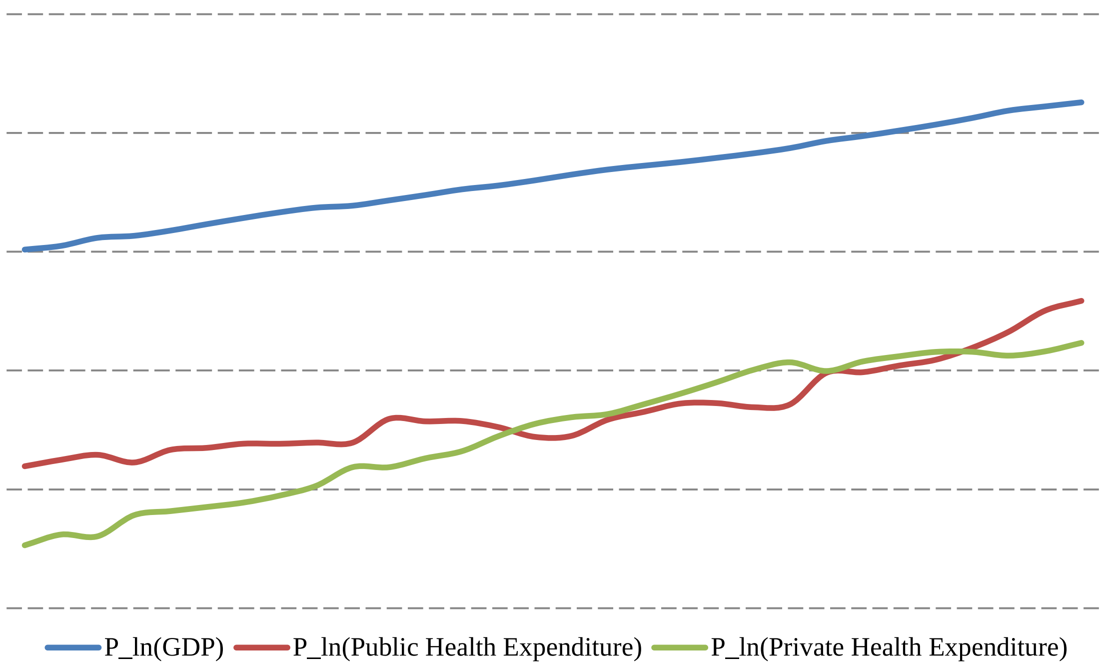


Figure 4.6. Permanent components by multivariate BN approach.

For the cyclical components, several characteristics can be observed from Figure 4.7. First, public health expenditure is pro-cyclical. The correlation between GDP and the public health expenditure cyclical series is 0.406 (significant at 5% level) (see Table 4.5 for details). This finding is consistent with the literature that states that while social spending is usually acyclical in developed countries, it is more likely to be pro-cyclical in developing countries (Arze del Granado *et al*., 2010). Next, observe that while there is a negative correlation between GDP and the private health expenditure cyclical series of -0.253, this is not significant, consistent with a view that private health expenditure is acyclical. The third observation is that the cyclical components of health expenditure are more volatile than for GDP (see summary statistics in Table 4.1 for details).

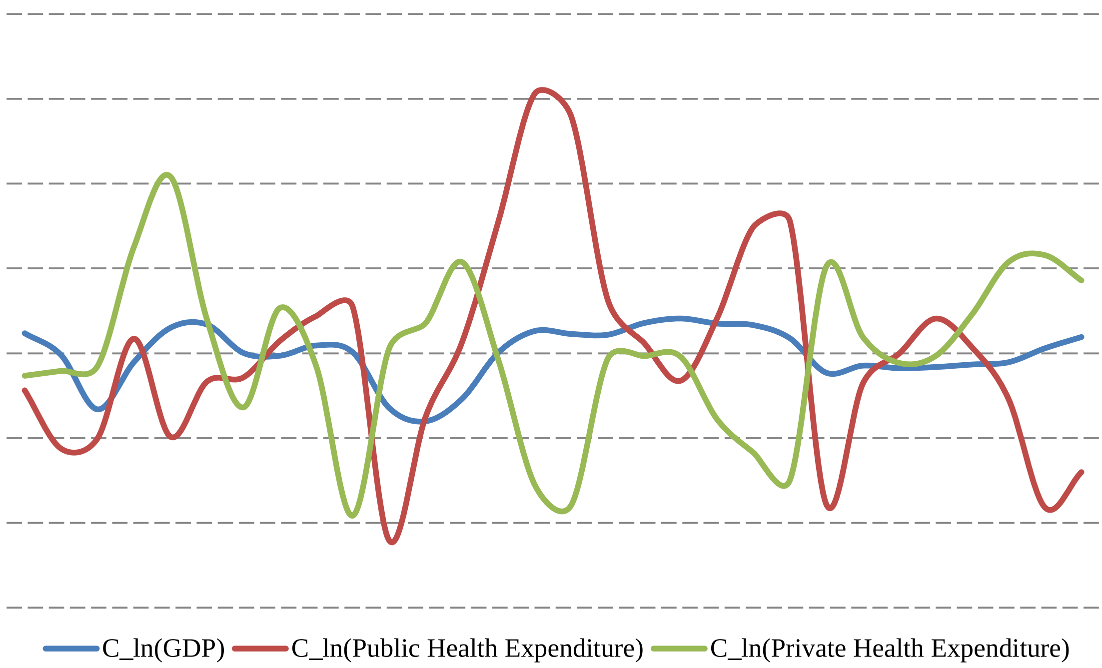


Figure 4.7. Transitory components by multivariate BN approach.

Table 4.5. Correlation statistics for the multivariate Beveridge-Nelson decomposed series.

Permanent components				Cyclical components			
	$P_{\ln(GDP)}$	$P_{\ln(PubHE)}$	$P_{\ln(PHE)}$		$C_{\ln(GDP)}$	$C_{\ln(PubHE)}$	$C_{\ln(PHE)}$
$P_{\ln(GDP)}$	1			$C_{\ln(GDP)}$	1		
$P_{\ln(PubHE)}$	0.941***	1		$C_{\ln(PubHE)}$	0.406**	1	
$P_{\ln(PHE)}$	0.980***	0.868***	1	$C_{\ln(PHE)}$	-0.253	-0.669***	1

Note: *, **, and *** indicate the Sidak-adjusted significance level at 10, 5, and 1% respectively.

The pro-cyclical public health expenditure and acyclical private health expenditure behaviours presented here are purely correlation relationships. The results may be due to omitted variable bias or bias caused by reverse causality. The next section will study the above relationship in more detail using provincial-level panel data.

4.5. REGIONAL PANEL DATA ANALYSIS RESULTS

4.5.1 The Cyclical Behaviour of Government Health Expenditure

The simple correlation statistics of the variables used in the regression analysis are presented in Panel A of Table 4.6. There are positive correlations between government health expenditure growth and economic growth (r = 0.444), and between government health expenditure growth and central government transfer growth (r = 0.185). On the other hand, there is a negative relationship between economic growth and central government transfer growth (r = -0.128). The correlations between other control variables in Panel A and government health expenditure are all quite low. The regression results using this group of variables are reported in Columns (2) and (3) in each table.

In Panel B, the first difference terms of three control variables (the percentage of population aged between 0 and 14 (POP014), the percentage of population aged 65 and older (POP65P), and urbanisation rates (URB)) are included as the original levels. The correlations between POP014, POP65P, URB and government health expenditure growth are stronger than their counterparts in Panel A. Except for the POP014 variable, all the other control variables show positive correlation with dependent variables. The regression results adopting this specification are reported in Columns (4) and (5) in the following tables.

Table 4.6. Simple correlation statistics of variables in the government health expenditure equation.

Panel A							
	Δln(GHE)	Δln(GRP)	Δln(TRANSFER)	Δ(CDR)	Δ(POP014)	Δ(POP65P)	Δ(URB)
Δln(GHE)	1						
Δln(GRP)	0.444	1					
Δln(TRANSFER)	0.185	-0.128	1				
Δ(CDR)	0.088	0.022	0.094	1			
Δ(POP014)	0.040	-0.017	0.027	0.098	1		
Δ(POP65P)	0.041	-0.017	0.038	0.064	0.038	1	
Δ(URB)	0.033	0.115	-0.053	0.153	-0.032	-0.035	1
Panel B							
	Δln(GHE)	Δln(GRP)	Δln(TRANSFER)	Δ(CDR)	POP014	POP65P	URB
POP014	-0.312	-0.435	0.125	-0.056	1		
POP65P	0.306	0.403	-0.128	0.048	-0.758	1	
URB	0.138	0.260	-0.215	0.037	-0.795	0.580	1

The heteroskedasticity and autocorrelation tests results are reported in Table 4.7. As can be seen, the null hypothesis of homoscedasticity is rejected at 1% level for all equations; the null of no serial correlation cannot be rejected at 10% level for all equations. Based on above discussion, the heteroskedastic robust standard error is used. For the FE estimator, the bootstrap heteroskedasticity robust standard error is chosen; for the IV estimator, the GMM estimator plus a heteroskedasticity robust standard error is adopted.

Table 4.7. Heteroskedasticity and autocorrelation tests results for government health expenditure equations.

	(1)	(2)	(3)	(4)	(5)
Panel A: whole China					
Heteroskedasticity test†	182.45***	229.64***	215.76***	113.56***	108.55***
Autocorrelation test‡	0.104	0.024	0.044	0.309	0.357
Panel B: coastal China					
Heteroskedasticity test†	57.39***	60.86***	61.08***	191.48***	204.07***
Autocorrelation test‡	0.687	0.658	0.961	0.431	0.687
Panel C: inland China					
Heteroskedasticity test†	61.88***	74.53***	80.24***	39.64***	36.38***
Autocorrelation test‡	1.155	0.239	0.550	1.375	1.953

Notes: †Modified wald statistic for groupwise heteroskedasticity in fixed effect model with a null hypothesis of homoskedasticity. ‡Wooldridge test for serial correlation with a null hypothesis of no first-order autocorrelation. *** p<0.01, ** p<0.05, * p<0.1. Full control variables included and the no. of observations please refer to Tables 4.8-4.10.

The fixed-effects estimates for the government health expenditure equation are reported in Tables 4.8 to 4.10, using whole, coastal and inland China samples respectively. Five columns of results are presented in each table. Except for the time dummy for the year 2003 in Column (1), only the key control variable, Δln(GRP) is included. The following two columns further include the Δln(TRANSFER), Δ(CDR) and a subgroup of *x* vector while the first difference terms (of POP014, POP65P and URB) are adopted. The last two columns include the subgroup of *x* vector as their original terms for robustness analysis.

Starting from the results using the whole China sample in Table 4.8, pro-cyclical behaviour of government health spending can be observed, as the coefficients of the provincial-level economic growth variable are robustly significant and positive. Depending on which other covariates are included, the magnitude ranges from 1.034 to 3.793. The second key variable of interest is central government transfer growth. The significantly positive coefficients (ranging from 0.128 to 0.253) in Columns (2) to (5) suggest that central government transfers are also a key factor influencing the provincial-level governments' health spending behaviour.

Two health related variables, the change in the crude death rate and the time dummy for the year 2003 are consistently insignificant. For the sub-group of variables in the *x* vector, which are firstly included according to the empirical model specification (Equation 4.3) as first-difference terms, the coefficients are all insignificant (as shown in Columns (2) and (3)); the effects of the changes of age structure (the proportion of younger or older population) and urbanisation rates are minor. If those three variables are included as levels, as shown in Columns (4) and (5), they tend to all be significant. The proportion of the population that is young has a negative association with government expenditure growth, while an ageing population has a much stronger positive association with government expenditure growth. The latter finding is potentially due to the fact that government health expenditure is responding to the increased demand for health that goes with the ageing population. The former finding is unusual, and it would be better to include dummies for early childhood (i.e. 5 years old or younger), and the middle childhood and early adolescence years (i.e. aged between 6 and 14) separately. An increasing proportion of the population in early childhood may have a positive correlation with government spending through its role as public health provider (e.g. immunization programs), while the reasons for the increasing share of middle childhood and early adolescence are more unclear. Data limitations impede any further investigation here.

Finally, urbanisation is associated with a higher growth of government health expenditure. This may be due to the disparity of government expenditure in urban and rural areas through its funding in public hospitals. As mentioned in Chan *et al*. (2008, page 349), of the entire government subsidy to health care in 2002, 67.7% went to hospitals; however, there are huge urban-rural disparities where 50.5% was allocated to urban hospitals while only 7.3% was allocated to health centres (which are the main health services providers) in rural areas.

Splitting the sample into coastal and inland China, several changes are observed (see Tables 4.9 and 4.10). Firstly, although there is some evidence to suggest that although the coefficients of economic growth variables are all positive, they are no longer consistently significant. As can be found in Table 4.9 for the coastal China sample and Table 4.10 for the inland China sample, when POP14, POP65P and URB variables are included as levels, the economic growth variable is no longer significant in both samples; the central government transfer growth in coastal China sample is also insignificant. This finding suggests heterogeneity between the two major regions.

Another change worth mentioning here is the change of urbanisation in the coastal China sample. For example, in Columns (2) and (3) of the coastal China sample (Table 4.9), the change in the urbanisation rate shows a significant negative association with government health expenditure growth, while in Columns (4) and (5) the level of urbanisation is no longer significant. This finding suggests that in coastal China, where the urbanisation rate is relatively high, there may exist non-linear associations between urbanisation and government health spending behaviour.

To summarise, the FE estimates provide robust evidence suggesting government health expenditure is exhibiting pro-cyclical behaviour in the full sample, and the growth of central government transfers has also shown a positive association with the growth of government health spending. As mentioned in the methodology section, the FE estimates may be subject to endogeneity bias. As such, the results using IV/GMM strategies are now discussed next.

Table 4.8. Cyclicality of government health expenditure: fixed-effects estimates (whole China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	3.639***	3.793***	3.788***	1.034**	1.080**
	[0.477]	[0.478]	[0.486]	[0.516]	[0.517]
Δln(TRANSFER)		0.245***	0.253***	0.128**	0.134**
		[0.051]	[0.052]	[0.054]	[0.055]
Δ(CDR)			0.017		0.007
			[0.014]		[0.014]
Δ(POP014)		0.287	0.196		
		[0.810]	[0.822]		
Δ(POP65P)		0.983	0.688		
		[0.889]	[0.800]		
Δ(URB)		-0.168	-0.148		
		[0.293]	[0.285]		
POP014				-0.694*	-0.710*
				[0.405]	[0.401]
POP65P				4.069***	3.977***
				[0.957]	[0.944]
URB				0.444**	0.435**
				[0.208]	[0.207]
D2003	-0.006	0.015	0.014	0.006	0.006
	[0.015]	[0.018]	[0.018]	[0.016]	[0.017]
No. of Provinces	31	31	31	31	31
No. of Observations	401	401	399	401	399

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Bootstrap standard errors in brackets (the bootstrap is based on 1000 redrawings and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

Table 4.9. Cyclicality of government health expenditure: fixed-effects estimates (coastal China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	2.817***	2.941***	2.923***	0.635	0.591
	[0.583]	[0.541]	[0.558]	[0.708]	[0.718]
Δln(TRANSFER)		0.108*	0.106*	-0.005	-0.009
		[0.056]	[0.056]	[0.051]	[0.052]
Δ(CDR)			0.016		0.018
			[0.016]		[0.014]
Δ(POP014)		1.832	1.833		
		[1.416]	[1.423]		
Δ(POP65P)		0.153	0.136		
		[0.903]	[0.906]		
Δ(URB)		-1.157***	-1.195***		
		[0.386]	[0.369]		
POP014				-0.857	-0.862
				[0.955]	[0.957]
POP65P				2.709**	2.792**
				[1.350]	[1.388]
URB				0.350	0.342
				[0.438]	[0.442]
D2003	0.033	0.045	0.044	0.026	0.025
	[0.022]	[0.027]	[0.028]	[0.025]	[0.026]
No. of Provinces	12	12	12	12	12
No. of Observations	156	156	156	156	156

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Bootstrap standard errors in brackets (the bootstrap is based on 1000 redrawings and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

Table 4.10. Cyclicality of government health expenditure: fixed-effects estimates (inland China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	4.026***	4.282***	4.293***	1.178	1.249
	[0.723]	[0.739]	[0.753]	[0.778]	[0.782]
Δln(TRANSFER)		0.340***	0.354***	0.242***	0.260***
		[0.083]	[0.089]	[0.080]	[0.086]
Δ(CDR)			0.016		-0.005
			[0.022]		[0.024]
Δ(POP014)		-0.028	-0.163		
		[1.027]	[1.072]		
Δ(POP65P)		1.380	0.530		
		[1.746]	[1.506]		
Δ(URB)		0.125	0.183		
		[0.297]	[0.304]		
POP014				-0.705*	-0.709*
				[0.397]	[0.402]
POP65P				4.532***	4.439***
				[1.282]	[1.288]
URB				0.561**	0.559**
				[0.228]	[0.226]
D2003	-0.028*	0.022	0.021	0.017	0.020
	[0.016]	[0.031]	[0.034]	[0.027]	[0.029]
No. of Provinces	12	12	12	12	12
No. of Observations	245	245	243	245	243

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Bootstrap standard errors in brackets (the bootstrap is based on 1000 redrawings and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

The results using an instrumental variable approach are reported in the following tables. Firstly, classical instrument variables used in applied macro analysis are used, in which the current value is instrumented by the lagged value. Following this logic, the current economic growth and change in crude death rate are instrumented by their one-year lagged values. The results using this strategy are reported in Tables 4.11 to 4.13. Secondly, for robustness analyses an additional instrumental variable, the yearly temperature difference, is added and the key results are reported in Table 4.14. In all the results tables the weak identification test - Kleibergen-Paap *rk* Wald *F* statistic (Kleibergen and Paap, 2006) assuming non independent, identically distributed residuals - is reported. Rejecting the null hypothesis suggests that the instruments are not weak. As suggested by Baum *et al*. (2007), the critical values complied by Stock and Yogo (2005) for the i.i.d. case or the “rule of thumb” that the *F* statistic should be at least 10 (Staiger and Stock, 1997) could be used. Since the Staiger and Stock’s rule of thumb corresponds to a 5% level test that the maximum size is no more than 15% when the number of instruments is one or two (Stock and Yogo, 2005), the critical value of 15% maximum IV size is used in this thesis to infer that whether the instruments are weak or not. According to this rule, except for two indicated equations in Table 4.14, the instruments are found to be not weak. The second test statistic reported in the results table is a test of overidentifying restrictions when there are more instruments than endogenous variables. The Hansen *J* statistics are reported and the results suggest the null hypothesis of valid instruments cannot be rejected in all equations. The last test statistic included in the results table is an endogeneity test used to test whether an endogenous variable can be treated as exogenous. Rejecting the null hypothesis suggests that IV estimates are preferred.

Firstly, the first stage result of Column (3), Panel A, in Table 4.14 is used as an example to show that both the lagged value and weather shock variable are valid instruments. The Kleibergen-Paap *rk* Wald *F* statistic and Hansen *J* statistics are already reported in the table and thus not replicated here. The first stage estimated result for Δln(GRP) equation is:

$$\Delta \ln(GRP_{it}) = 0.760 \times L.\Delta \ln(GRP_{it}) - 0.000775 \times TempDiff_{it} + z_{1it}$$
$$\qquad [0.0428]^{***} \qquad\qquad [0.000339]^{**}$$

where the robust standard errors are reported in the brackets. The coefficient of the lagged economic growth is positive and significant at 1%, the coefficient of the weather shock is negative and significant at 5%, *L* is the lag operator. The detailed results of other variables included in the z_1 vector (Δln(TRANSFER), Δ(POP014), Δ(POP65P), Δ(URB), *L*.Δ(CDR) and D2003) are not reported for brevity. As anticipated by theory, the weather shock has a significant and negative effect on economic growth.

The first-stage estimated result for Δ(CDR) equation is:

$$\Delta \ln(MOR_{it}) = -0.369 \times L.\Delta(CDR_{it}) - 0.00327 \times TempDiff_{it} + z_{2it}$$
$$\qquad [0.0670]^{***} \qquad\qquad [0.00804]$$

where the robust standard errors are reported in brackets, the coefficient of the lagged crude death rate change is negative and significant at 1%, the coefficient of the weather shock is insignificant at 10%, *L* is the lag operator, the detailed results of other variables included in the z_2 vector (Δln(TRANSFER), Δ(POP014), Δ(POP65P), Δ(URB), *L*.Δln(GRP) and D2003) are not reported for simplicity.

There may be concerns about whether the previous year's economic growth may also have an impact on government health expenditure. For example, people may have been over-worked during economic expansion in the last year, so their health status in the next year could be influenced and their health expenditure could consequently increase.

Increased health expenditure is more likely to come from the individuals as OOP expenditure or from medical insurance expenditure. Although not perfect, the change in the crude death rate variable may capture some of the unobservable changing health status effects. A natural way to do this is to use the two period lagged values as the instrumental variable. The results based on this strategy are comparable with results reported here; however, there are small first stage *F* statistics and a large magnitude for both estimated coefficients and standard errors. For example, if using the two year lagged value as the instrumental variable, the estimated coefficient of Δln(GRP) and Δln(TRANSFER) are 7.349 and 0.298, corresponding to an estimated coefficient of 4.986 and 0.265 in Column (3), Table 4.10 (all are significant by 1% criteria). Based on the above consideration, the classical one-year lagged value would be a better choice in this static panel analysis.

As can be seen from Table 4.11, the magnitudes of the coefficients of Δln(GRP) variable in all columns become larger compared with their counterparts reported in Table 4.8. Moreover, the coefficients are all significantly positive. As for the subsample analyses, the results found economic growth is significant and positive for the coastal China sample (Table 4.12). Without including the insignificant Δ(CDR) variables, all other columns also suggest a similar conclusion for the inland China sample (Table 4.13); government health expenditure is pro-cyclical. However, the endogeneity test results in Columns (4) & (5) suggest that in these two equations, the endogenous variables could actually be regarded as exogenous. In this case, the corresponding results reported in Table 4.10 are preferred.

The coefficient on the growth in central government transfers is positive and significant for the full and inland China samples, while, depending on the control variables included, the effects are either insignificant or significantly positive in the coastal China sample. Usually omitted in cross-country analyses, the growth in central government transfers deserves more attention in this country specific analysis. As mentioned earlier, this variable is used to capture the redistribution of tax revenue from the central government. From this perspective the results reported so far firstly show that provinces in inland China rely more on central government transfers to fund local government health expenditure, than those in coastal China. Secondly, since the central government transfer relies on tax revenue, a significantly positive coefficient on central government transfer growth may represent a general pro-cyclical behavior with regard to economic growth of a whole country.

Finally, several robustness checks have been undertaken in this static panel setting. Firstly, adding weather shocks as an additional instrumental variable does not change the main results; see Table 4.14. Secondly, adding common time dummies for each year to capture potential unobservable shocks further reinforces the significant and positive coefficient on the economic growth variables (now the coefficients are also robustly significant in the inland China sample). However, the coefficients on central government transfers become insignificant; see Tables 4.15 and 4.16 for details.

Table 4.11. Cyclicality of government health expenditure: IV/GMM estimates (whole China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	4.726***	5.036***	4.986***	2.020***	1.698**
	[0.451]	[0.449]	[0.486]	[0.677]	[0.778]
Δln(TRANSFER)		0.265***	0.265***	0.157***	0.132**
		[0.051]	[0.052]	[0.049]	[0.054]
Δ(CDR)			0.035		0.079
			[0.060]		[0.058]
Δ(POP014)		0.233	0.116		
		[0.586]	[0.681]		
Δ(POP65P)		1.042	0.638		
		[0.803]	[0.777]		
Δ(URB)		-0.290	-0.331		
		[0.308]	[0.317]		
POP014				-0.554	-0.704
				[0.440]	[0.492]
POP65P				3.250***	3.312***
				[0.964]	[0.982]
URB				0.416**	0.366**
				[0.164]	[0.179]
D2003	-0.011	0.012	0.010	0.006	-0.001
	[0.016]	[0.017]	[0.018]	[0.016]	[0.019]
Weak identification test statistic	324.645	308.019	14.886	65.559	16.033
Endogeneity test statistics	13.516	18.496	18.047	3.324	5.460
No. of Provinces	31	31	31	31	31
No. of Observations	401	401	398	401	398

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Heteroskedasticity robust standard errors in brackets. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. The weak identification test statistics suggest that the instruments are not weak; all p values for endogeneity test statistics are lower than 0.10.

Table 4.12. Cyclicality of government health expenditure: IV/GMM estimates (coastal China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	4.396***	4.453***	4.451***	2.335**	2.299**
	[0.657]	[0.620]	[0.640]	[0.931]	[0.970]
Δln(TRANSFER)		0.111*	0.110*	0.047	0.042
		[0.060]	[0.058]	[0.059]	[0.061]
Δ(CDR)			0.003		0.023
			[0.091]		[0.082]
Δ(POP014)		1.702*	1.702*		
		[0.911]	[0.910]		
Δ(POP65P)		0.400	0.397		
		[0.819]	[0.818]		
Δ(URB)		-1.401***	-1.409***		
		[0.413]	[0.432]		
POP014				-0.546	-0.549
				[0.676]	[0.675]
POP65P				1.455	1.545
				[1.272]	[1.383]
URB				0.337	0.327
				[0.284]	[0.293]
D2003	0.022	0.035	0.034	0.025	0.023
	[0.026]	[0.027]	[0.027]	[0.026]	[0.027]
Weak identification test statistics	103.438	50.917	6.275	31.905	6.809
Endogeneity test statistics	9.462	10.735	11.808	4.916	5.759
No. of Provinces	12	12	12	12	12
No. of Observations	156	156	156	156	156

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Heteroskedasticity robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. The weak identification test statistics suggest that the instruments are not weak; all p values for endogeneity test statistics are lower than 0.10.

Table 4.13. Cyclicality of government health expenditure: IV/GMM estimates (inland China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	4.879***	5.307***	5.174***	1.721**	1.018
	[0.595]	[0.596]	[0.666]	[0.868]	[1.084]
Δln(TRANSFER)		0.367***	0.360***	0.257***	0.190**
		[0.085]	[0.095]	[0.077]	[0.092]
Δ(CDR)			0.055		0.120
			[0.076]		[0.075]
Δ(POP014)		-0.029	-0.328		
		[0.743]	[0.984]		
Δ(POP65P)		1.306	0.027		
		[1.501]	[1.434]		
Δ(URB)		0.060	-0.010		
		[0.347]	[0.429]		
POP014				-0.634	-1.009
				[0.589]	[0.751]
POP65P				4.072***	3.873**
				[1.421]	[1.532]
URB				0.538***	0.482**
				[0.193]	[0.220]
D2003	-0.031	0.023	0.016	0.018	-0.003
	[0.019]	[0.029]	[0.033]	[0.025]	[0.031]
Weak identification test statistics	221.429	203.511	8.706	38.047	10.498
Endogeneity test statistics	5.662	8.535	7.922	0.596†	3.968†
No. of Provinces	19	19	19	19	19
No. of Observations	245	245	242	245	242

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Heteroskedasticity robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. The weak identification test statistics suggest that the instruments are not weak. Unless otherwise indicated, all p values for endogeneity test statistics are lower than 0.10. † p > 0.10.

Table 4.14. Cyclicality of government health expenditure: IV/GMM estimates, robust analyses by adding temperature yearly difference as an additional instrument variable.

	(1)	(2)	(3)	(4)	(5)
Panel A: whole China					
Δln(GRP)	4.727***	5.060***	5.005***	1.985***	1.658**
	[0.450]	[0.447]	[0.484]	[0.670]	[0.772]
Δln(TRANSFER)		0.278***	0.278***	0.159***	0.133**
		[0.050]	[0.051]	[0.049]	[0.054]
Δ(CDR)			0.038		0.080
			[0.060]		[0.057]
Weak identification test statistics	189.569	181.897	9.904	37.465	10.750
Overidentification test statistics	0.356	1.248	1.255	0.184	0.254
Endogeneity test statistics	13.162	17.446	17.018	3.148	5.223
Panel B: coastal China					
Δln(GRP)	4.231***	4.378***	4.372***	2.031**	2.006**
	[0.642]	[0.614]	[0.634]	[0.890]	[0.932]
Δln(TRANSFER)		0.136**	0.128**	0.047	0.040
		[0.057]	[0.057]	[0.058]	[0.060]
Δ(CDR)			0.053		0.047
			[0.086]		[0.079]
Weak identification test statistics	89.452	89.767	11.362	21.748	4.682†
Overidentification test statistics	1.792	2.891	3.086	1.989	2.371
Endogeneity test statistics	7.637	8.374	9.927	3.432	4.492‡
Panel C: inland China					
Δln(GRP)	4.846***	5.287***	5.143***	1.754**	1.018
	[0.593]	[0.587]	[0.651]	[0.874]	[1.091]
Δln(TRANSFER)		0.364***	0.354***	0.251***	0.191**
		[0.082]	[0.091]	[0.077]	[0.093]
Δ(CDR)			0.057		0.123
			[0.075]		[0.075]
Weak identification test statistics	113.980	105.622	5.803†	20.097	7.127
Overidentification test statistics	0.271	0.036	0.046	1.254	0.894
Endogeneity test statistics	5.708	8.538	7.957	0.724‡	4.795

Notes: Dependent variable: Change in Log Real Government Health Expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Heteroskedasticity robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. Only the coefficients of Δln(GRP), Δln(TRANSFER), and Δln(CDR) are reported here, full control variables included and the no. of observations please refer to Table 4.11-4.13. Unless otherwise indicated, the weak identification test statistics suggest that the instruments are not weak; † F statistics lower than 20% maximal IV size. All p values for overidentification statistics are higher than 0.05. Unless otherwise indicated, all p values for endogeneity test statistics are lower than 0.10; ‡ p > 0.10.

Table 4.15. Cyclicality of government health expenditure: fixed-effects estimates, robust analyses by adding common time dummies.

	(1)	(2)	(3)	(4)	(5)
Panel A: whole China					
Δln(GRP)	1.776***	1.796***	1.835***	1.512***	1.551***
	[0.435]	[0.428]	[0.427]	[0.434]	[0.431]
Δln(TRANSFER)		0.076*	0.086**	0.065	0.073*
		[0.042]	[0.038]	[0.041]	[0.038]
Δ(CDR)			0.016		0.016
			[0.011]		[0.011]
Panel B: coastal China					
Δln(GRP)	2.125***	2.082***	2.058***	1.895***	1.844**
	[0.701]	[0.723]	[0.739]	[0.714]	[0.723]
Δln(TRANSFER)		0.032	0.025	0.025	0.017
		[0.057]	[0.055]	[0.056]	[0.054]
Δ(CDR)			0.015		0.014
			[0.012]		[0.013]
Panel C: inland China					
Δln(GRP)	1.431**	1.492***	1.602***	1.349**	1.458**
	[0.582]	[0.578]	[0.587]	[0.612]	[0.614]
Δln(TRANSFER)		0.066	0.084	0.062	0.084
		[0.062]	[0.062]	[0.062]	[0.063]
Δ(CDR)			0.019		0.016
			[0.016]		[0.017]

Notes: Dependent variable: Change in Log Real Government Health Expenditure (Δln(GHE)). Bootstrap standard errors in brackets (the bootstrap is based on 1000 redrawings and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1. Only the coefficients of Δln(GRP), Δln(TRANSFER), and Δln(CDR) are reported here, full control variables included and the no. of observations please refer to Table 4.8-4.10.

Table 4.16. Cyclicality of government health expenditure: IV/GMM estimates, robust analyses by adding common time dummies.

	(1)	(2)	(3)	(4)	(5)
Panel A: whole China					
Δln(GRP)	2.168***	2.289***	2.334***	2.053***	2.063***
	[0.686]	[0.682]	[0.684]	[0.702]	[0.712]
Δln(TRANSFER)		0.079*	0.082*	0.069	0.063
		[0.046]	[0.045]	[0.045]	[0.045]
Δ(CDR)			0.030		0.043
			[0.043]		[0.043]
Weak identification test statistics	73.115	71.587	14.892	65.709	16.853
Endogeneity test statistics	0.639	1.032	1.137	1.016	1.429
Panel B: coastal China					
Δln(GRP)	2.906***	2.879***	2.884***	2.960***	2.946***
	[1.026]	[1.004]	[0.993]	[1.010]	[1.020]
Δln(TRANSFER)		0.030	0.026	0.028	0.019
		[0.063]	[0.062]	[0.064]	[0.067]
Δ(CDR)			0.008		0.016
			[0.070]		[0.069]
Weak identification test statistics	69.045	65.788	5.586	47.565	6.358
Endogeneity test statistics	1.192	1.311	1.456	1.777	2.025
Panel C: inland China					
Δln(GRP)	1.722**	1.768**	1.770**	1.781**	1.762**
	[0.827]	[0.818]	[0.860]	[0.823]	[0.879]
Δln(TRANSFER)		0.069	0.042	0.066	0.035
		[0.066]	[0.070]	[0.065]	[0.070]
Δ(CDR)			0.080		0.087
			[0.056]		[0.053]
Weak identification test statistics	27.940	27.216	7.837	27.783	10.289
Endogeneity test statistics	0.211	0.188	1.295	0.446	2.071

Notes: Dependent variable: Change in Log Real Government Health Expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Heteroskedasticity robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. Only the coefficients of Δln(GRP), Δln(TRANSFER), and Δln(CDR) are reported here, full control variables included and the no. of observations please refer to Table 4.11-4.13. The weak identification test statistics suggest that the instruments are not weak. All p values for endogeneity test statistics are higher than 0.10.

Further considering the potential lagged effects of fiscal policy, lagged dependent variables are included. First, the one period lagged dependent variable is considered. If the regression residuals cannot pass the Lagrange Multiplier test for residual serial correlation, the two period lagged dependent variable is also included. The Arellano-Bover/Blundell-Bond estimates for the whole, coastal and inland China samples are reported from Tables 4.17 to 4.19.

Starting with Table 4.17, significant pro-cyclical behaviour for government health expenditure can be observed for the whole China sample. Central government transfer growth is also robustly significant and positive. When splitting the sample into two groups, the results are mixed. Firstly, the lagged dependent variable is found to be insignificant when other covariates are included (see Table 4.18 for details). This may cast doubt on whether it is appropriate to use dynamic panel frameworks to analyse the coastal China sample. As for the inland China sample, the coefficient of economic growth is still positive, but now mixed as to whether it is significant or not. The central government transfer growth variable is again robustly significant and positive. Finally, as shown in the robustness analysis, when adding additional instrumental variables, the conclusions remain the same. See Table 4.20 for details.

Table 4.17. Cyclicality of government health expenditure: Arellano-Bover / Blundell-Bond GMM estimates (whole China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	1.816***	2.273***	2.487***	1.497***	1.857***
	[0.295]	[0.530]	[0.590]	[0.370]	[0.613]
Δln(TRANSFER)		0.221***	0.274***	0.228***	0.254***
		[0.074]	[0.086]	[0.070]	[0.091]
Δ(CDR)			0.056*		0.056
			[0.031]		[0.040]
Δ(POP014)		0.846	0.238		
		[0.891]	[0.908]		
Δ(POP65P)		1.022	0.542		
		[0.770]	[1.513]		
Δ(URB)		-0.569	-0.794		
		[0.692]	[0.486]		
POP014				0.506	0.967
				[0.700]	[1.016]
POP65P				2.748**	2.911**
				[1.263]	[1.295]
URB				0.253	0.197
				[0.231]	[0.324]
L.Δln(GHE)	0.186***	0.147*	0.089	0.081	0.117
	[0.070]	[0.079]	[0.076]	[0.082]	[0.088]
L2.Δln(GHE)	0.474***	0.470***	0.461***	0.413***	0.480***
	[0.051]	[0.061]	[0.070]	[0.072]	[0.079]
AR(1)	0.002	0.000	0.000	0.001	0.001
AR(2)	0.594	0.470	0.511	0.812	0.236
Sargan test statistic	30.316	29.751	28.855	30.332	27.885
No. of Provinces	31	31	31	31	31
No. of Observations	339	339	337	339	337

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Windmeijer WC-robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. All p values for Sargan test statistic are higher than 0.1. AR(1) & AR(2) are the p values for the first and second order Lagrange Multiplier test for residual serial correlation.

Table 4.18. Cyclicality of government health expenditure: Arellano-Bover / Blundell-Bond GMM estimates (coastal China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	2.153**	2.169*	4.302	1.021	3.018
	[0.980]	[1.240]	[8.793]	[1.340]	[2.409]
Δln(TRANSFER)		-0.009	0.146	0.037	0.154
		[0.123]	[0.423]	[0.153]	[0.164]
Δ(CDR)			0.049		-0.064
			[0.187]		[0.209]
Δ(POP014)		0.884	-0.693		
		[2.115]	[8.334]		
Δ(POP65P)		4.168	0.963		
		[3.768]	[15.104]		
Δ(URB)		-0.763	2.687		
		[1.353]	[13.724]		
POP014				-1.119	5.034
				[8.975]	[4.216]
POP65P				4.349	-6.809
				[9.827]	[10.951]
URB				0.248	2.412
				[2.777]	[2.588]
L.Δln(GHE)	0.437***	0.278	0.039	0.109	0.361
	[0.142]	[0.238]	[1.180]	[0.311]	[0.529]
L2.Δln(GHE)					0.568
					[0.431]
AR(1)	0.034	0.063	0.383	0.113	0.291
AR(2)	0.120	0.281	0.693	0.228	0.514
Sargan test statistic	11.845	7.849	8.723	9.417	6.654
No. of Provinces	12	12	12	12	12
No. of Observations	144	144	144	144	132

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Windmeijer WC-robust standard errors in brackets. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. All p values for Sargan test statistic are higher than 0.1. AR(1) & AR(2) are the p values for the first and second order Lagrange Multiplier test for residual serial correlation.

Table 4.19. Cyclicality of government health expenditure: Arellano-Bover / Blundell-Bond GMM estimates (inland China).

	(1)	(2)	(3)	(4)	(5)
Δln(GRP)	1.761***	2.180*	2.917*	0.748	2.260
	[0.609]	[1.172]	[1.584]	[1.025]	[2.082]
Δln(TRANSFER)		0.329***	0.341**	0.306***	0.364***
		[0.091]	[0.160]	[0.093]	[0.094]
Δ(CDR)			0.035		0.047
			[0.050]		[0.064]
Δ(POP014)		1.793	-0.144		
		[1.953]	[1.179]		
Δ(POP65P)		-3.562	0.365		
		[7.377]	[3.644]		
Δ(URB)		0.069	-0.415		
		[0.639]	[2.510]		
POP014				1.887*	1.791
				[1.038]	[1.465]
POP65P				6.585***	4.988
				[1.956]	[3.293]
URB				0.749*	0.317
				[0.386]	[0.583]
L.Δln(GHE)	0.100	0.052	-0.018	-0.029	-0.058
	[0.070]	[0.106]	[0.109]	[0.070]	[0.118]
L2.Δln(GHE)	0.473***	0.456***	0.459***	0.417***	0.441***
	[0.056]	[0.087]	[0.131]	[0.100]	[0.111]
AR(1)	0.010	0.003	0.010	0.005	0.006
AR(2)	0.592	0.435	0.543	0.327	0.237
Sargan test statistic	18.838	17.220	17.340	17.692	16.204
No. of Provinces	19	19	19	19	19
No. of Observations	207	207	205	207	205

Notes: Dependent variable: change in natural log real government health expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Windmeijer WC-robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. All p values for Sargan test statistic are higher than 0.1. AR(1) & AR(2) are the p values for the first and second order Lagrange Multiplier test for residual serial correlation.

Table 4.20. Cyclicality of government health expenditure: Arellano-Bover / Blundell-Bond GMM estimates, robust analyses by adding temperature yearly difference as an additional instrument variable.

	(1)	(2)	(3)	(4)	(5)
Panel A: whole China					
Δln(GRP)	1.827***	2.272***	2.550***	1.593***	1.953**
	[0.301]	[0.538]	[0.582]	[0.588]	[0.762]
Δln(TRANSFER)		0.228***	0.270***	0.226***	0.268***
		[0.079]	[0.098]	[0.071]	[0.092]
Δ(CDR)			0.059*		0.035
			[0.031]		[0.030]
AR(1)	0.003	0.001	0.000	0.001	0.001
AR(2)	0.523	0.492	0.398	0.799	0.400
Sargan test statistic	30.707	29.557	30.050	30.582	29.542
Panel B: coastal China					
Δln(GRP)	2.122**	2.756***	-5.826	1.751	3.405
	[0.891]	[0.874]	[12.338]	[1.143]	[3.068]
Δln(TRANSFER)		-0.037	0.305	-0.013	0.038
		[0.126]	[0.608]	[0.163]	[0.127]
Δ(CDR)			0.078		0.029
			[0.150]		[0.221]
AR(1)	0.032	0.023	0.208	0.042	0.271
AR(2)	0.122	0.191	0.686	0.240	0.704
Sargan test statistic	11.864	7.575	5.304	7.520	4.790
Panel C: inland China					
Δln(GRP)	1.848***	2.433*	4.461*	1.001	2.662
	[0.583]	[1.246]	[2.502]	[1.030]	[1.888]
Δln(TRANSFER)		0.300***	0.373**	0.311***	0.335***
		[0.097]	[0.156]	[0.088]	[0.126]
Δ(CDR)			0.024		0.053
			[0.035]		[0.131]
AR(1)	0.010	0.004	0.077	0.004	0.007
AR(2)	0.568	0.488	0.814	0.389	0.270
Sargan test statistic	18.790	18.202	16.014	17.708	16.966

Notes: Dependent variable: Change in Log Real Government Health Expenditure (Δln(GHE)). Instrumented variable: Δln(GRP) & Δ(CDR). Windmeijer WC-robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. Only the coefficients of Δln(GRP), Δln(TRANSFER), and Δln(CDR) are reported here, full control variables included and the no. of observations please refer to Table 4.17-4.19. All p values for Hansen J statistic are higher than 0.05. AR(1) & AR(2) are the p values for the first and second order Lagrange Multiplier test for residual serial correlation.

4.5.2 The Cyclical Behaviour of Private Health Expenditure

The simple correlation statistics for variables used in the private health expenditure equation are presented in Table 4.21. As can be seen, income and basic medical insurance expenditure variables have a positive correlation with private health expenditure, while government health expenditure has a fairly weak and negative relationship with the dependent variable. The heteroskedasticity and autocorrelation tests results are reported in Table 4.22. As can be seen, the null hypothesis of homoscedasticity is rejected at the 1% level for all equations; the null of no serial correlation cannot be rejected at the 10% level for first seven columns. Similar to the government health expenditure equations, the bootstrap standard error is chosen for the FE estimator.

The fixed-effects estimates of the cyclicality of per capita private health expenditure are reported in Table 4.23. As discussed in the methodology section there may be reverse causality between private health expenditure and disposable income, and between private health expenditure and mortality. Since the lagged value is not a valid instrument here, as suggested by fairly low first stage F statistics, only results from FE estimators without using an IV approach are reported. Strictly speaking, the causality cannot be inferred and coefficients reported should be viewed as correlations in the Private Health Expenditure equations.

Three columns of regression results have been reported for each sample area (whole, coastal, and inland China). In the first column, except for the time dummy for the year

2003, only two key control variables are included, the first difference of natural log of per capita real disposable income of urban residents (hereafter the disposable income growth), and the first difference of the natural log of per capita real urban basic medical insurance expenditure (hereafter the basic medical insurance expenditure growth). In the second and third columns, a similar group of control variables to those used in the cyclicality of government health expenditure equations have been included. The central government transfer is not included here since it will not have a direct effect on private spending behaviour, and its indirect effect will be captured by the newly added government health expenditure variable.

Several conclusions can be drawn. Firstly, significantly positive coefficients on per capita disposable income growth are found across all nine columns, suggesting a potential for pro-cyclical behaviour for private health expenditure. Secondly, the coefficients on per capita basic medical insurance expenditure growth are also positive and significant. If the expansion of basic medical insurance has an effective financial protection effect on the individual, the coefficient should be negative, so a positive estimate suggests that the financial protection effect of basic medical insurance is quite limited. However, to investigate further the effect of medical insurance, micro-level survey data would be useful.

When a variety of control variables are included in the model, the coefficients change only slightly, so the conclusions remain the same. The main clearly significant effect in the control variables is the variable capturing proportion of the population aged under 15, which has a strong positive effect for the full sample and for inland China. This is not the case for the government health expenditure equations: it appears the extra demand for health services induced through younger children is mainly financed through private means, not government support. It is also worth noting that the coefficients of government health expenditure growth and crude death rate growth are consistently insignificant.

Finally, in another robustness analysis, the dummy variable for the year 2003 has been replaced by a set of time dummies in each year. The pro-cyclical behaviour of private health expenditure is still present for the whole China sample and inland sample, but not for the coastal China sample, although it is still positive. The coefficient on the basic medical expenditure growth is no longer significant, suggesting the financial effect of medical insurance is still limited.

Table 4.21. Simple correlation statistics of variables in the private health expenditure equation.

Panel A								
	Δln(PHE)	Δln(INCO)	Δln(BMI)	Δln(GHE)	Δ(CDR)	Δ(POP014)	Δ(POP65P)	Δ(URB)
Δln(PHE)	1							
Δln(INCO)	0.332	1						
Δln(BMI)	0.329	0.117	1					
Δln(GHE)	-0.013	0.221	-0.211	1				
Δ(CDR)	0.098	0.082	0.052	0.023	1			
Δ(POP014)	0.098	-0.012	-0.182	0.079	0.191	1		
Δ(POP65P)	0.077	0.077	0.126	-0.002	0.063	0.020	1	
Δ(URB)	0.050	0.026	-0.145	-0.057	0.140	0.105	-0.070	1
Panel B								
	Δln(PHE)	Δln(INCO)	Δln(BMI)	Δln(GHE)	Δ(CDR)	POP014	POP65P	URB
POP014	0.038	-0.169	0.131	-0.116	0.035	1		
POP65P	-0.036	0.154	-0.177	0.167	0.015	-0.697	1	
URB	-0.022	0.155	-0.088	-0.005	-0.012	-0.813	0.584	1

Table 4.22. Heteroskedasticity and autocorrelation tests results for private health expenditure equations.

	whole China			coastal China			inland China		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Heteroskedasticity test†	872.75***	777.53***	473.68***	119.07***	135.96***	488.02***	445.16***	448.92***	176.04***
Autocorrelation test‡	1.330	0.161	0.115	0.738	0.006	0.199	0.686	4.298*	4.677**

Notes: †Modified wald statistic for groupwise heteroskedasticity in fixed effect model with a null hypothesis of homoskedasticity. ‡Wooldridge test for serial correlation with a null hypothesis of no first-order autocorrelation. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Full control variables included and the no. of observations please refer to Table 4.23.

Table 4.23. Cyclicality of per capita private health expenditure of urban residents: fixed-effects estimates.

	whole China			coastal China			inland China		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Δln(INCO)	1.057***	1.081***	0.906***	1.121**	1.266**	1.063**	1.016***	0.963***	0.878**
	[0.270]	[0.292]	[0.314]	[0.564]	[0.517]	[0.478]	[0.325]	[0.340]	[0.421]
Δln(BMI)	0.119***	0.129***	0.092***	0.125***	0.126**	0.094*	0.119***	0.136***	0.102**
	[0.029]	[0.029]	[0.030]	[0.047]	[0.049]	[0.050]	[0.045]	[0.041]	[0.046]
Δln(GHE)		-0.030	0.106		-0.157	0.055		0.017	0.095
		[0.058]	[0.086]		[0.147]	[0.169]		[0.052]	[0.098]
Δ(CDR)		0.003	0.023		0.035	0.011		-0.007	0.021
		[0.038]	[0.040]		[0.090]	[0.083]		[0.041]	[0.046]
Δ(POP014)		2.038***			0.484			2.856***	
		[0.678]			[2.203]			[0.604]	
Δ(POP65P)		0.146			-1.226			0.691	
		[1.265]			[1.695]			[1.795]	
Δ(URB)		0.470			-0.236			1.262	
		[0.719]			[1.148]			[0.919]	
POP014			0.824			0.439			1.043
			[0.685]			[1.853]			[0.691]
POP65P			-0.858			-1.633			-1.548
			[0.954]			[1.611]			[1.728]
URB			-0.573*			-0.877			0.099
			[0.340]			[0.625]			[0.475]
D2003	-0.015	-0.017	-0.044*	0.021	0.033	-0.019	-0.039*	-0.035	-0.056**
	[0.023]	[0.025]	[0.025]	[0.046]	[0.051]	[0.056]	[0.021]	[0.021]	[0.025]
No. of Provinces	31	31	31	12	12	12	19	19	19
No. of Observations	247	247	247	96	96	96	151	151	151

Notes: Dependent variable: change in natural log real per capita private health expenditure of urban residents (Δln(PHE)). Bootstrap standard errors in brackets (the bootstrap is based on 1000 redrawings and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

Table 4.24. Cyclicality of per capita private health expenditure of urban residents: fixed-effects estimates, robust analyses by adding common time dummies.

	whole China			coastal China			inland China		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Δln(INCO)	0.996***	0.922***	0.937***	0.930*	0.901	0.929	1.132***	1.067***	1.097***
	[0.330]	[0.343]	[0.331]	[0.547]	[0.631]	[0.582]	[0.392]	[0.372]	[0.365]
Δln(BMI)	0.029	0.031	0.029	0.050	0.054	0.051	0.014	0.008	0.012
	[0.049]	[0.048]	[0.049]	[0.108]	[0.121]	[0.126]	[0.058]	[0.054]	[0.060]
Δln(GHE)		0.089	0.081		0.045	0.042		0.104	0.100
		[0.096]	[0.104]		[0.184]	[0.169]		[0.096]	[0.106]
Δ(CDR)		-0.001	0.007		-0.043	-0.038		-0.004	0.006
		[0.032]	[0.035]		[0.090]	[0.091]		[0.031]	[0.035]
No. of Provinces	31	31	31	12	12	12	19	19	19
No. of Observations	247	247	247	96	96	96	151	151	151

Notes: Dependent variable: change in natural log real per capita private health expenditure of urban residents (Δln(PHE)). Bootstrap standard errors in brackets (the bootstrap is based on 1000 redrawings and the replications is based on provinces clusters). *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Full control variables included please refer to Table 4.23.

4.6. DISCUSSION AND CONCLUSIONS

This chapter has investigated the cyclical behaviour of public and private health expenditure in China, using both national time series and provincial-level panel data. Correlation analysis based on decomposition of the aggregate time series suggests that at a national level, public health expenditure tends to be pro-cyclical while private health expenditure tends to be acyclical. To gain further insight into these relationships, provincial-level panel data were utilised. Adopting several panel data modeling techniques, results confirm this finding that government health expenditure is indeed pro-cyclical, and there is strong evidence for a pro-cyclical causal link from economic growth to government health expenditure. As for private health expenditure, there is some evidence which suggests that it is also pro-cyclical with regard to disposable income in urban China. This is not inconsistent with the national-level findings, where no significant correlation was found, leading to the tentative conclusion of acyclical behavior. With this richer, province-level data set, the evidence is clearer for pro-cyclical private health expenditure. However, there is not sufficient information in the data to be able to say whether there is a causal relationship between growth and private expenditure.

Two caveats are worth mentioning here and may be investigated in the future when more data become available. Firstly, for the national time series decomposition analysis, the accuracy of the results may be impacted by the relatively short time span (32 years). However these are the only available data relevant to this topic in China. Nonetheless, the results found are consistent with health care policy transition in China. Specifically, the findings from the permanent components of the series identify two main health financing strategy transitions in the past three decades. The first transition, around 1993, is in accordance with decentralisation when the government shifted the main financial burden to the population. The second transition occurs after the outbreak of SARS when the government realised the importance of health to the economy and decided to undertake the key role of health care financing again. Secondly, for the time period investigated, despite the recent global financial crisis, China has still maintained a positive economic growth rate. So strictly speaking, the analysis is unable to speak to the likely cyclical behaviour of health expenditure during severe recessions. However, there is still a great deal of variation in the potential long run growth rate that can help identify cyclical behaviours. The results found in this paper are comparable with studies using cross-country samples by Arze del Granado *et al*. (2010), as presented in Table 4.25

Table 4.25. Comparison of estimated cyclical coefficients of public health expenditure.

Sources	Region		
Panel A: World sample			
Arze del Granado (2010)	Average	Good times	Bad times
	1.25**	1.51**	0.56
Panel B: sub-sample analyses by countries groups			
Arze del Granado (2010)	Higher middle income	Middle income	Low income
	1.75**	1.64**	3.02**
Panel C: China sample			
Chapter 4	Whole	coastal	inland
Time dummy for year 2003†	2.020***	2.335***	1.721*
Time dummy for all years‡	2.053***	2.960***	1.781*

Notes: †The results are drawn from Table 4.11-4.13; ‡the results are drawn from Table 4.16; since the change of crude death rate are robust insignificant, estimated coefficient from columns (4) in each table are selected and presented here. *** p<0.01, ** p<0.05, * p<0.1.

Owing to the relatively short-time horizon, one sub-analysis this chapter does not perform is to separate the good times and bad times[45] as done by Arze del Granado (2010). This analysis may be worth proceeding with when a longer time series is available. Whether the pro-cyclical behaviour of health expenditure is symmetric or asymmetric deserves more attention, especially from a policy-makers' perspective. If the pro-cyclical effect is symmetric for the good and bad times, then economic downturns (especially economic crises) could have profound impacts on the health expenditure plans of national governments and the ability to pay for health care of individuals. More importantly, unless a well-developed social medical insurance scheme is in place, the poor and the vulnerable (e.g. women, children, elderly, informal sector workers, migrant workers, and unemployed) could be the first to experience negative health-related consequences (Gottret *et al*., 2009; Parry and Humphreys, 2009; Semenza *et al*., 2012). More detailed policy suggestions for protecting pro-poor health services during economic crises are discussed in Gottret *et al*. (2009).

To minimize the potential negative effect of economic downturns on population health in China in the future, two policy suggestions are highlighted based on lessons learnt from past economic crises.

(1) Universal medical insurance schemes should be established and proved to show a financial protection effect, as they have been found to effectively protect health service utilization during economic crises.

[45] The good times and bad times are defined as whether the output gaps (observed real series minus the potential permanent series decomposed by the Hodrick-Prescott filter) are positive or negative (Arze del Granado *et al*., 2010).

A comparative analysis between Thailand and Indonesia during the Asian financial crisis suggests that a well-developed medical insurance scheme in Thailand had a significant effect on the cyclical behavior of health during economic downturns (Waters *et al.*, 2003). As shown in O'Donnell *et al*. (2008), China, Nepal, and Bangladesh are three Asian countries for which OOP health expenditure accounted for more than 60% of total health expenditure at the beginning of the new century. When the major responsibility of health care finance lies with individuals, the poor and vulnerable may face a higher risk of unaffordable health care expense during economic downturns. The Chinese government has responded to this issue by gradually expanding voluntary basic medical insurance coverage since 2003. However, as discussed in this paper and reported in several empirical analyses using survey data, the poor and vulnerable were more likely to be excluded from this social protection program for three reasons. First, taking the recently established URBMI as an example, being non-eligible to purchase this basic medical insurance is a fundamental barrier for the migrant workers who do not have local urban *hukou* (type of household registration) (Lin *et al*., 2009). Secondly, although the government has implemented a special policy for subsidizing poor local households to purchase URBMI, as found in Chen and Yan (2012), among urban residents with the lowest 25% of income in the survey, the people who did not qualify for the special government subsidy were significantly less likely to purchase voluntary basic medical insurance, suggesting ability to pay is still a barrier for universal medical insurance coverage. Thirdly, the current basic medical insurance schemes have not established a solid financial protection effect. Wagstaff *et al*. (2009a) found that although NRCMS had improved the health service utilization rates, the OOP inpatient expenditure per visit had also significantly increased. Universal coverage of medical insurance plus a risk-based benefit package could provide effective protection of health care accessible for the population, especially for the poor, during economic downturns.

(2) The limited government health budgets should be used efficiently, and there is a need to set priorities for health investment strategies during economic downturns.

Health care resources are limited, and not all services can be provided to the whole population. There is a need to optimize health benefits and reduce health inequality based on a scarce health budget through priority setting. Economic evaluation (e.g. cost-effectiveness analysis) is a powerful tool that can be used for priority setting (Hauck *et al*., 2004). If government health expenditure has to be cut, then governments could subsidize and protect cost-effective public health and essential medicines during an economic crisis.

Take a key component of current health reform in China – pharmaceutical reform – as an example. Pharmaceutical expenditure accounts for a large proportion of total health expenditure in China – 43% of an average inpatient episode and 51% of an outpatient visit (Sun *et al*., 2008). As discussed in the introduction, as a consequence of fiscal decentralization since economic reform, health care providers have to fund this expense themselves. It is thus not surprising that selling drugs has become a major source of revenue for hospitals, and drug use fees have become the most profitable item (Meng *et al*., 2005). As a consequence, substantial overprescribing is widely observed (Li *et al*., 2012). Furthermore, the choice of drugs is sometimes made according to profit margins rather than clinical efficacy (Sun *et al*., 2008). Since pharmaceutical expenditures are not fully reimbursed by basic medical insurance, the potential OOP cost for pharmaceuticals could be a barrier to health services access during economic downturns when individuals' capability to fund health care decreases, let alone the poor and the vulnerable that might be excluded from any basic medical insurance. Containment of pharmaceutical expenditures should be a top priority for policy intervention. Encouraging generic drug competition and substitution, which has been adopted in some developed countries, could also be an effective policy intervention tool for pharmaceutical cost-containment in developing countries like China (Danzon and Chao, 2000; Wang, 2006; Cameron *et al*., 2012). A case study using public hospital data in China suggests that by switching four medicines from branded drugs to the lowest price generic drugs could save US$370 million (Cameron *et al*., 2012). Other intervention policies (e.g. reform of the fee-for-service payment system, drug lists and hospital revenue capping policies) have been discussed elsewhere in detail (Liu *et al*., 2000; Hu *et al*., 2001; Yip *et al*., 2012).

It is of concern that population health outcomes are subject to potential risks during economic downturns. The current findings will help policymakers in China and other developing countries to consider what fiscal policies might be adopted in case of recession. The next chapter will make use of publicly available aggregate health outcomes data to explore the relationship between economic growth, economic fluctuations and health outcomes.

CHAPTER 5

ECONOMIC DEVELOPMENT, FLUCTUATIONS AND HEALTH OUTCOMES IN CHINA

5.1. INTRODUCTION

Maternal and child mortality, malnutrition status and tuberculosis (TB) prevalence are among key indicators listed in the Millennium Development Goals that 189 nations are committed to improve by 2015 (United Nation Development Programme (UNDP), 2010). The majority of maternal and child deaths, child malnutrition and TB incidence occurs in developing countries, representing a huge disparity in health between the developed and developing worlds (Ruger and Kim, 2006; Olness and Kalanj, 2009).

Maternal and child mortality are regarded as headline indicators for measuring health, well-being and socio-economic conditions of populations. Reducing the maternal mortality ratio (MMR) has been a priority for developing countries (it accounts for 99% of deaths), especially in Sub-Saharan Africa and South Asia where 87% of global maternal deaths occur (World Health Organization (WHO) *et al*., 2010). A similar situation is found for the under-five mortality rate (U5MR). Around 50% of the world's under-five deaths in 2010 occurred in five developing countries: India, Nigeria, Democratic Republic of the Congo, Pakistan and China (Hill *et al*., 2012).

Child malnutrition is also a vital indicator as it is found to be significantly related to human capital accumulation and labour productivity in later adulthood. Low birth weight is also likely to decrease the survival chances for newborns (Marmot and Wadsworth, 1997; Victoria *et al.*, 2008; Spence and Lewis, 2009). The most vulnerable populations live in developing countries, and are usually the children of parents with the lowest social status. According to the United Nations (UN) (2011), Southern Asia, Sub-Saharan Africa and South-Eastern Asia are the regions with the highest proportion of children under age five who are underweight in 2009 (43%, 22% and 18%, respectively). Evidence also suggests that children from the poorest households or who live in rural areas are more likely to be underweight in Southern Asia (UN, 2011).

TB is still one of the biggest infectious killers, although the incidence rate has slowed down globally since 2004. In 2009, India, China, South Africa, Nigeria and Indonesia were the top five countries with the largest number of new TB cases, while India and China alone accounts for 35% of the world's share. Asia and Africa accounts for around 85% of the total of new TB cases (UN, 2011). TB can impact on the economy through its direct costs of illness to the individual and to the country, and the indirect cost of lost productivity (Jackson *et al*., 2006).

Health inequalities between developed and less developed countries in the world intuitively suggest that economic development may be a key factor that is positively related to the improvement of health outcomes. Indeed, empirical analyses show that the relationship is not only associative but causal (see Pritchett and Summers (1996) for an example using instrumental variable (IV) strategy, and Lindahl (2005) for an example investigating a truly exogenous part of income – monetary lottery prizes), i.e. the wealthier the healthier. The difficulty in drawing a causal relationship as to whether increased wealth/income can cause an improvement in health status is partly due to the existence of reverse causality, i.e. improved health status may also lead to an improvement in wealth. The literature which focuses on the impact of health outcomes/health investment on economic output was summarised and discussed in detail in Chapters 2 and 3. Another possibility for the potential endogeneity issue comes from the unobservable factors which may influence income and health simultaneously (e.g. genetics, personality traits; see Imlach Gunasekara *et al*. (2011) for a review). Unless the endogeneity issue has been carefully handled, the estimated coefficient should be examined with caution.

Micro data have been widely used in studying the relationship between income and health. Imlach Gunasekara *et al*. (2011) provides a detailed literature review of panel studies investigating whether changes in income impact on changes in self-reported health in adults. In most of studies they reviewed, increases in income have a small and positive impact on self-reported health. Regarding the relationship between household income and the health of children, there is evidence suggesting the presence of gender bias in household resource allocation in rural areas. Using survey data from Romania, Skoufias (1998) found that in rural areas, higher household income (proxied by monthly per capita expenditure) is significantly positively (negatively) associated with the short-run nutritional status (e.g. weight-for-height Z-scores) of boys (girls); this association is insignificant in urban areas. Regarding long-run nutritional status (health-for-age Z-scores), higher household income is found to be significantly positive only for girls in rural areas. Macro data have also been utilised to investigate the impact of income changes on health changes and similar positive correlation has been reported. Using cross-country data, Biggs *et al*. (2010) further provide evidence that the impact of income on health also depends on the distribution of income.

Another group of studies focuses on the relationship between economic fluctuation and health outcomes. As discussed in Chapter 4, there are several channels via which *short-run* economic fluctuations can impact on health outcomes such as the accessibility of quality health care (Waters *et al.*, 2003; Hopkins, 2006; Gottret *et al.*, 2009; Anderson *et al.*, 2011). Furthermore, within a single country the impact of economic fluctuations may be distributed unevenly, with a stronger effect on more vulnerable populations. For newborn infants born into low socio-economic status families, economic hardship can put them at greater risk of illness with less chance of survival in the first year (Lin, 2006). The dominant role of out-of-pocket (OOP) health expenditure in health care financing[46] in less developed countries may warrant special attention during economic fluctuations (perhaps especially during economic downturns).

[46] Take Asian countries/regions as an example, O'Donnell *et al*. (2008) shows that compared to Japan (in which OOP accounts for 12.8% of total health expenditure), in all other 12 countries/regions' OOP health expenditure accounts for more than 30% of total health care finance in the late 1990s and 2000. OOP health expenditure accounts for even more than 60% of total health expenditure in three countries: Nepal (75.0%), Bangladesh (64.6%) and China (60.4%).

Economic fluctuations may impact on health outcomes via two dimensions: transitory and permanent effects. Thus far the evidence is inconclusive regarding both these effects. Micro-level survey data are important resources with which to study whether there is a permanent effect of economic fluctuations on health outcomes. Using a natural experiment in France, Banerjee *et al*. (2010) estimates the long-run health effect of income shocks experienced in early childhood. The authors found that the effect of negative income shocks varies with regard to health measures: while there was significant negative impact on height by age 20, there was no impact on infant mortality, life expectancy, and morbidity by age 20. On the other hand, there is experimental evidence from a cash transfer program suggesting that positive income shocks can benefit the mental health of school-age girls; however, the effect was only limited to the intervention period (Baird *et al*., 2012).

Macro data have been widely used in this area to study the effect of economic fluctuations on health outcomes from a transitory perspective. Although it is hypothesised that there is a *pro*-cyclical pattern of health outcomes and economic fluctuations (i.e. health outcomes become worse during economic downturns, or vice versa) as discussed above, the direction reported in the literature is actually inconclusive, depending on the chosen health outcomes and the regions of interest (for example, poor countries versus rich countries), see Suhrcke and Stuckler (2012) for a detailed review.

Studies analysing general age-adjusted mortality rates, using data from the US, suggest that mortality moves pro-cyclically with economic expansion (Tapia Granados, 2005). In other words, recessions are ‘good’ for population’ health. In a subsequent paper using data from

Japan, Tapia Granados (2008) shows that while general mortality and age-specific mortality show a similar *pro*-cyclical pattern, economic fluctuations have a different effect on age-specific mortality rates (i.e. *counter*-cyclical at ages 20-44, while *pro*-cyclical at ages 45-64 and ages 65-84 groups) or disease-specific mortality rates. Evidence from Canada generally supports the above findings. Ariizumi and Schirle (2012) found that *pro*-cyclical patterns are strongest in the 30-39 age group. A limitation of these earlier studies is that often the conclusion is attained from simply regressing health outcomes (or the deviation of health outcomes) on economic fluctuation indicators while excluding other relevant variables, i.e. they suffer from omitted variable bias. This makes it difficult to judge whether the results would still hold if other covariates were included. For example, as reported in Ariizumi and Schirle (2012), when excluding other relevant control variables, unemployment is significant at the 10% level; however the coefficient became insignificant when other covariates were included.

For the key indicators of interest in this study, maternal and child health outcomes (e.g. maternal, infant, neonatal mortality rates and low birth weight), it is hypothesised that economic fluctuations, especially recessions and debt crises, will hinder improvements in maternal and child health outcomes as can be seen from the Child Survival Initiative and the Safe Motherhood Initiative in developing countries in the late 1970s and 1980s (Price, 1994; Geefhuysen, 1999; Langer *et al*., 1999). However, more recent empirical studies have found mixed results. Evidence from Taiwan (Lin, 2006) and Argentina (Cruces *et al*., 2012) found a *counter*-cyclical movement with economic expansion. By analysing German state data (1980-2000), Neumayer (2004) found infant mortality and neonatal mortality is acyclical if the unemployment rate is used, while the above two child health indicators move pro-cyclically if the change in real Gross Domestic Product (GDP) is adopted. Schady and Smitz (2010), using data from 17 middle and low income countries, found that in most countries, infant mortality rates tend to move *pro*-cyclically or are acyclical with GDP; only when faced with very deep economic shocks is the counter-cyclical relationship found. Using the data around the 1997-98 Asian financial crisis period from Indonesia, Waters *et al.* (2003) found that the economic crisis had led to an increase in morbidity levels; however, the deterioration in nutritional status was minor. Stuckler *et al.* (2009), using data from 26 European Union countries, (1970-2007), shows that maternal mortality and infant mortality are acyclical with the increased unemployment rate. The results are robust even with a massive increase in unemployment.

The empirical evidence from regression analyses regarding infectious diseases (e.g. TB in this study) prevalence is limited. Rechel *et al*. (2011) conducted a qualitative analysis regarding the global economic crisis which began in 2008, and communicable disease control among experts from 23 European countries. The key result revealed the concern that prevention services are susceptible to budget cuts during an economic crisis; the economic hardship may further increase the pool of the vulnerable population and increase the risk of infectious disease. Suhrcke *et al*. (2011) conducted a detailed systematic review of evidence (in which 76% were descriptive analyses) related to the impact of economic crises on communicable disease transition and control. Among the 37 studies they reviewed, 30 studies identified an adverse infectious disease outcome along with the economic crises; the other 7 studies found a positive and/or insignificant outcome. The evidence reviewed by Suhrcke *et al*. suggests that the effect of an economic crisis depends critically upon budgetary responses by government.

The literature reviewed above suggests that there is a heterogeneous effect of economic development/fluctuations on health outcomes across countries. The differences in health systems, health care finance and delivery may be one relevant factor. From this perspective, to inform country specific policies for China, a country-specific study is required. This chapter contributes to the literature by further providing empirical evidence using data from mainland China.

Currently, knowledge of the relationship between economic development/fluctuation and health outcomes in China from a macro perspective is limited. Banister and Zhang (2005), adopting three years of provincial-level panel data (1985, 1990 and 1995), found that there is a significant and positive relationship between economic development (as proxied by per capita consumption) and improved health outcomes (child, adult mortality and life expectancy). Wang *et al*. (2006) adopted longer provincial-level panel data and found some evidence suggesting that living standards (as proxied by expenditure on food consumption as a percentage of total income) are significant and positively associated with maternal mortality (while insignificantly associated with crude death rates). Uchimura and Jütting (2009) and Jin and Sun (2011) found economic development (proxied by natural logarithm of GRP per capita) is negatively associated with infant mortality rates. Apart from Jin and Sun (2011), the other studies did not discuss the potential endogeneity issue of the income variable. Without providing details, Jin and Sun suggested that the null hypothesis (that income is exogenous) cannot be rejected. There is one study investigating the relationship between economic cycle and mortality in China using national-level time series data. Song *et al*. (2011) studies the nexus between economic cycles and occupational accidents and they found no significant relationship during the pre-reform period (1953-1978), while a significant pro-cyclical relationship (i.e. occupational fatalities increase with economic booms and decrease with economic depressions) during the reform period (1978-2008).

This chapter departs from the above literature in three main ways. Firstly, a wider selection of health outcome measures is considered. The current literature mostly focuses on child or maternal mortality indicators; little is known about the relationship between economic development and nutrition indicators and/or infectious diseases from a macro perspective. Secondly, the potential endogeneity issue of economic development (or income) variables, which are seldom discussed in studies related to China, is formally tested and dealt with. Thirdly, the relationship between economic fluctuations and health outcomes is investigated. The results may not only be relevant for China but also for other developing countries in transition.

5.2. METHODOLOGY AND DATA

5.2.1. Economic and Econometric Framework

In this chapter, both national time series data and provincial-level panel data are used to investigate the relationship between economic growth and health outcomes.

Because only 19 years of time series data are available at a national level, the first set of results involve a group of simple bivariate correlation analyses between economic development/fluctuations proxies and health outcomes are performed. Since separate city (hereafter urban) and rural level mortality data are available, the natural logarithm of per capita real annual disposable income of urban household (ln(DINCO_U)) or per capita real annual net income of rural household (ln(NINCO_R)) [47] is used to measure the economic development for urban areas and rural areas separately. To plot the correlation between economic fluctuations and health outcomes, the first difference in the natural logarithm of per capita real annual disposable income of urban households (Δln(DINCO_U)) or per capita real annual net income of rural households (Δln(NINCO_R)) is used.

By relying on a richer provincial level panel dataset, the framework for analysing health outcomes is specified as:

$$HEALTH_{it} = \alpha_1 + \beta_1 ECO_{it} + \gamma_1' x_{it} + \mu_{1i} + \varepsilon_{1it} \tag{5.1}$$

where $HEALTH_{it}$ is a health outcome indicator of interest, ECO_{it} is the key independent variables capturing economic development, x_{it} is a vector that contains other relevant control variables based on previous literature and the availability of data in China, i indexes the province and t indexes the year, μ_{1i} is the time-invariant unobservable heterogeneity, ε_{1it} is an idiosyncratic error term. Assume that μ_{1i} is uncorrelated with ε_{1it}, and may be correlated with the other regressors in Equation (5.1). To include the potential region effect into the model, the model includes an interaction term between a region dummy (which is set to be 1 if for inland China and 0 otherwise) and economic development variable.

In the provincial-level analysis where only aggregated mortality data are available,

[47] According to the National Bureau of Statistics of China, the formula for calculating net income is "net income = total income - household operation expenses - taxes and fees-depreciation of fixed assets for production - gifts to rural relatives".

economic development is mainly measured as a natural logarithm of real per capita Gross Regional Product (GRP). The control variables included in the *x* vector are as follows. Firstly, the provincial government health expenditure[48] as a percentage of total provincial government expenditure (GHE_GOV) is used to measure the level of general government health investment, instead of using another common proxy, the government health expenditure per capita. As discussed in Chapter 4, government health expenditure is moving pro-cyclically with GDP; the level of government health expenditure is also determined by GDP. Therefore using a GHE per capita variable will absorb the effect of GDP on health outcomes. As shown in the literature (Gavin and Perotti, 1997; Lane, 2003; Talvi and Vegh, 2005), government expenditure in developing countries also moves pro-cyclically with GDP. In this case using the share of government health expenditure as a percentage of government expenditure can eliminate the fiscal expansion effect to some extent (since both numerator and denominator move pro-cyclically with GDP) while still capturing the trend of health investment in provincial governments. Whenever possible, a more outcome-specific health service access variable is used to investigate the direct effect of health service access. For example, in the MMR equation, the modern birth attendance (or skilled birth attendance) rate is included, suggesting that it is an effective way of reducing MMR (Qiu *et al.*, 2010; Green *et al.*, 2011). Secondly, a group of health related variables are considered: the percentage of population aged between 0 and 14 (POP014), the percentage of population aged 65 and older (POP65P), and time dummy for the year 2003 (D2003) (which is included to capture the national Severe Acute Respiratory Syndrome (SARS) shock). Thirdly, the urbanisation rate (URB) and the average schooling years per labor (ASY) are also included.

[48] This variable is a much more accurate measurement of health care resources compared to Wang *et al*. (2006), in which owing to the data availability, the local government's expenditure on cultural, science, education and health is used to proxy for the health care resources.

To investigate the potential impact of economic fluctuations on health outcomes the following equation is adopted from Neumayer (2004) and Lin (2006).

$$HEALTH_{it} = \alpha_2 + \beta_2 \Delta(ECO_{it}) + \beta_3 ECO_{it} + \gamma_2' x_{it} + \mu_{2i} + \varepsilon_{2it} \tag{5.2}$$

where Δ is the first difference operator. The definition of other variables is the same as used in Equation (5.1).

The key indicator of economic fluctuations is the change of natural log of real GRP per capita (Δln(GRP per capita)). Another popular indicator for the measurement of economic instability is the unemployment rate. However using the unemployment rate is not suitable in this study owing to the availability and quality of data. Currently only the registered unemployment data for urban areas are available from the China Labour Statistical Yearbook. By area, the total employed persons could be divided into urban and rural employees in mainland China. In 2009 urban employees accounted for around 40% of total employed persons. Thus using the unemployment rate of urban employees alone will not capture the economic fluctuations in rural areas.

A fixed-effects (FE) estimator is adopted to estimate Equations (5.1) and (5.2) so that the unobservable province-specific time-invariant variables (as captured by μ_i) which are potentially correlated with both income and health (e.g. institution) will not be a concern. Omitted variable bias can be eliminated to some extent through using the FE estimator; however, the endogeneity issue cannot be ruled out as two key explanatory variables in this study - economic development and government health expenditure - are potentially endogenous. Unless the endogeneity issue is carefully handled, the estimated coefficient should be interpreted as associative, rather than causal. For the economic development variable, potential reverse causality is the main concern, as previously discussed in Chapters 2 and 3 - health capital can have a positive impact on economic output. Whether reverse causality is severe or not may depend on the health outcomes of interest. Pritchett and Summers (1996) pointed out that one benefit of studying child health is that there could be less concerns about potential reverse causation problems compared to the analyses of adult health and income. With regard to the infectious disease status, on the other hand, it could be closely associated with work productivity and thus reverse causality is hard to ignore. To handle the potential endogeneity issue, fixed assets investment (INV) is adopted as the instrument for GDP. It is hypothesised that fixed assets investment is positively correlated with GDP and it will not directly impact on individual health outcomes.

The potential endogeneity of government health expenditure is based on the view that local government responds to poor health outcomes in prior years by adjusting health spending in the current year. Finding a suitable instrument for government health expenditure is not an easy task. Bokhari *et al*. (2007) used the military expenditure of neighbouring countries as an IV for government health expenditure; however similar logic cannot be adopted in this provincial-level analysis. In this case, if a negative estimate of government health expenditure is expected, then the magnitude of true value will be larger than the estimates (Bokhari *et al*., 2007).

Finally, two key assumptions of the FE regression errors are tested. The first one is that the variances of the errors across provinces are identical (i.e. homoskedasticity). The assumption may not hold since in this analysis, the cross-sectional units have different size (i.e. population). The assumption is tested using a modified Wald statistic for groupwise heteroskedasticity in the residuals of a fixed effect regression model proposed by Greene (2002) and coded as “xttest3” in Stata software. The null hypothesis of homoscedasticity is rejected at 1% level for all equations. The second assumption to be tested is that conditional on individual effects μ_i, the error term should be serially uncorrelated. A Wald test proposed by Wooldridge (2002) and coded as “xtserial” in Stata (Drukker, 2003) is used. The null of no serial correlation is rejected at 1% level for all equations. To cope with the above two issues, standard errors that are robust to heteroskedasticity and autocorrelation is used. For the FE estimator, the bootstrap standard error is chosen; for the IV estimator, the Generalised Method of Moments (GMM) estimator plus a heteroskedasticity and autocorrelation consistent (HAC) robust standard error is adopted. To calculating the HAC robust standard error, the quadratic spectral (QS) kernel is used, and the bandwidth is set to be of order $T^{1/3}$ ($= 8^{1/3} \approx 2$) (Sun and Phillips, 2008).

To summarise, in the regression analysis for provincial-level panel data section, the following strategies will be used:

- FE estimates of economic development on health outcomes
- GMM/IV estimates of economic development on health outcomes
- FE estimates of economic fluctuations on health outcomes
- Using interaction terms between region dummy and economic development and/or economic fluctuation proxy to test the potential region difference

5.2.2. Data

Two different datasets (national time series and provincial-level panel data) are used in this chapter. It should be noted that, based on availability, different health outcomes and economic development measurements are used for the national time series analyses and the provincial-level panel analyses as mentioned above. This will be highlighted in due course.

5.2.2.1. National time series data

Four maternal and child mortality indicators are available for the national analysis at both rural and urban levels: MMR, U5MR, infant mortality rate (IMR), neonatal mortality rate (NMR). Mortality data is drawn from the *China Health Statistical Yearbook 2010*; other variables are drawn from the *China Statistical Yearbook*.

5.2.2.2. Provincial-level panel data

Firstly, four maternal and child health outcome measures are adopted for the provincial panel analysis: MMR, liveborn infants that weigh less than 2500g (LBW), perinatal mortality rate (PNM), and undernourishment rates among children under 5 years of age (MAL). The provincial-level maternal and child health data are gathered from the *China Health Statistical Yearbook* over several years. In 1989, the Chinese National Maternal and Child Health Surveillance System was established; within the system districts/counties that aimed to be nationally representative were selected based on the residence area (urban/rural), geographical location and socioeconomic status (Gao *et al*., 2009; Liang *et al*., 2010). Series quality control procedures during each level of data collection have been implemented to guarantee the data quality (see details at: Liang *et al*., 2010; Liu *et al*., 2011; Yi *et al*., 2011). Furthermore, another two infectious disease prevalence measures are also considered: TB mortality rate and TB incidence rate. The infectious disease data are compiled by the National Notifiable Infectious Disease Report System and were drawn from the *China Health Statistical Yearbook* over several years (see details at: Zhang and Wilson, 2012).

Other explanatory variables are defined as follows. Real GRP is calculated by combining the nominal value of GRP in 1978 and the GRP index. Real fixed assets investment is adjusted by the fixed assets investment index. The per capita form of the above variables are adjusted by the total population at year-end. The percentage of population aged between 0 and 14, aged 65 and older, has been included to control for differences in age structure. The urbanisation rate is defined as the percentage of the total population who live in the urban area. Average schooling years per worker is defined in Appendix of Chapter 4. The provincial government health expenditure and provincial government expenditure data are drawn from the *Finance Yearbook of China* over several years. The health services utilization data are drawn from the *China Health Statistical Yearbook* over several years. Unless specified, all provincial-level data comes from the *China Statistical Yearbook* in each year. The variable definition and summary statistics for national and provincial-level data are presented in Tables 5.1 and 5.2, respectively.

Table 5.1. Summary statistics for national data.

Variable	Unit	Mean	Std. Dev.	Min	Max	Obs.
MMR, city	deaths per 100,000 live births	31.700	7.421	22.300	46.300	19
MMR, county	deaths per 100,000 live births	67.679	19.424	34.000	100.000	19
U5MR, city	deaths per 1,000 live births	14.395	4.008	7.600	20.900	19
U5MR, rural	deaths per 1,000 live births	42.332	15.614	21.100	71.100	19
IMR, city	deaths per 1,000 live births	12.163	3.541	6.200	18.400	19
IMR, rural	deaths per 1,000 live births	34.489	12.331	17.000	58.000	19
NMR, city	deaths per 1,000 live births	9.500	2.708	4.500	13.900	19
NMR, rural	deaths per 1,000 live births	23.800	8.586	10.800	37.900	19
ln(DINCO_U)		7.555	0.441	6.886	8.325	19
ln(NINCO_R)		6.588	0.343	6.039	7.188	19
Δln(DINCO_U)		0.080	0.025	0.034	0.126	18
Δln(NINCO_R)		0.064	0.026	0.020	0.124	18

Table 5.2. Summary statistics for provincial-level data.

Variable	Unit	Mean	Std. Dev.	Min	Max	Obs
MMR	deaths per 100,000 live births	44.245	51.643	1.380	401.400	217
MMR, city	deaths per 100,000 live births	33.031	37.451	0.000	362.320	217
MMR, county	deaths per 100,000 live births	47.999	54.681	0.000	433.700	216
LBW	%	2.256	0.901	0.970	5.230	248
PNM	deaths per 1,000 live births	10.844	4.475	2.520	25.800	248
MAL	%	2.351	1.749	0.076	9.590	248
TB mortality rate	1/100,000	0.235	0.206	0.000	1.200	248
TB incidence rate	1/100,000	79.143	35.554	17.600	205.630	248
ln(GRP per capita)		8.296	0.632	6.863	10.278	248
Δln(GRP per capita)		0.110	0.023	0.030	0.213	248
ln(DINCO_U)		9.091	0.306	8.523	9.981	248
Δln(DINCO_U)		0.087	0.031	-0.072	0.190	248
ln(NINCO_R)		7.959	0.430	7.148	9.143	248
Δln(NINCO_R)		0.071	0.028	-0.012	0.160	248
ln(INV per capita)		8.749	0.652	7.327	10.364	248
GHE_GOV		0.048	0.012	0.027	0.077	248
Modern birth attendance rate	%	96.731	6.179	60.480	100.000	248
Modern birth attendance rate, city	%	98.172	3.676	72.990	100.000	248
Modern birth attendance rate, county	%	96.331	6.612	59.700	100.000	247
Antenatal examination rate	%	89.439	7.712	58.000	99.780	248
Neonatal supervision rate	%	83.743	10.367	39.522	97.680	248
POP014		0.189	0.046	0.076	0.283	248
POP65P		0.088	0.019	0.048	0.164	248
Urbanizarion rate		0.458	0.151	0.226	0.942	248
Average schooling years per worker		7.606	1.241	3.129	10.870	248

Note: The provincial level data for MMR in the year 2003 are missing.

5.3. RESULTS

5.3.1. Correlation Analysis from National Time Series Data

To understand the nature of the data a presentation of the bivariate relationships between the health outcomes and key independent variables (economic development/fluctuation) of interest in both rural and urban China is given by a series of correlation analyses and scatter plots.

5.3.1.1 Rural-Urban disparities in health outcomes

There is a decreasing trend in the MMR in both rural and urban areas over a 19 year timeframe, and the gap between rural and urban areas has also decreased. In 1991 the reported MMR is 100 deaths per 100,000 live births in rural areas, around 2.2 times higher than in urban areas (46.3 deaths per 100,000 live births). This gap has been narrowed down to a difference of 1.3 in 2009 (34 vs. 26.6 deaths per 100,000 live births). Although the current MMR is still higher than the European level (21 deaths per 100,000 live births in 2008), the figure is already far below the average level of South-East Asia (240 deaths per 100,000 live births in 2008) (WHO *et al*., 2010).[49]

A similar pattern is observed for three child mortality variables: U5MR, IMR and NMR. Firstly, the improvement of the U5MR is interesting, although the decline is a little slower in rural areas. The U5MR in rural areas has decreased from 71.1 to 21.1 deaths per 1,000 live births (3.4 vs. 2.8 times higher than in urban areas). However the mortality level in

[49] Unless specified, the regions (e.g. South-East Asia) refer to the WHO regions.

rural China in 2009 was still higher than the level in urban areas in 1991 (20.9 deaths per 1,000 live births). According to UN Inter-agency Group (2011), the U5MR of urban China in 2009 (7.6 deaths per 1,000 live births) has outperformed the level of Europe in 2010 (14 deaths per 1,000 livebirths), while the U5MR of rural China is also lower than the average level for South-East Asia (56 deaths per 1,000 live births).

Secondly, IMR in rural China has decreased from 58 deaths per 1,000 livebirths in 1991 to 17 deaths per 1,000 live births in 2009, attaining the level of urban areas in 1991 (17.3 deaths per 1,000 live births); the rural-urban difference has also decreased from 3.4 to 2.7 in the same period. The IMR of urban China in 2009 is better than the European level in 2010 (6.2 vs. 12 deaths per 1,000 live births), while the level of rural China is also much lower than the average South-East Asia level (44 deaths per 1,000 live births) (UN Inter-agency Group, 2011).

Finally, NMR has also improved considerably: it was 37.9 deaths per 1,000 live births in rural China in 1991 (3.0 times higher than urban area), and the number has decreased to 10.8 deaths per 1,000 live births (2.4 times higher than urban area) in 19 years. Comparing NMR globally, in 2009 the NMR in urban China is slightly higher than in the high income countries (4.5 vs. 3.6 deaths per 1,000 live births), while rural China is well below the South-East Asian level (30.7 deaths per 1,000 live births) and very close to the European level (10.7 deaths per 1,000 live births) (Oestergaard *et al*. 2011).

5.3.1.2. Bivariate correlation between economic development and health outcomes

The correlation between maternal and three child mortality indicators and economic development indicators (real personal income) in urban and rural area are presented in Figures 5.1-5.4. Within each figure, a correlation coefficient (r) is reported; furthermore, a straight line estimated from a simple bivariate regression is also presented. Although only 19 years' of time series data are available, it can still be clearly observed that an increased economic level is associated with a decreased mortality. There is some evidence that the correlation is stronger in rural areas than in urban China.

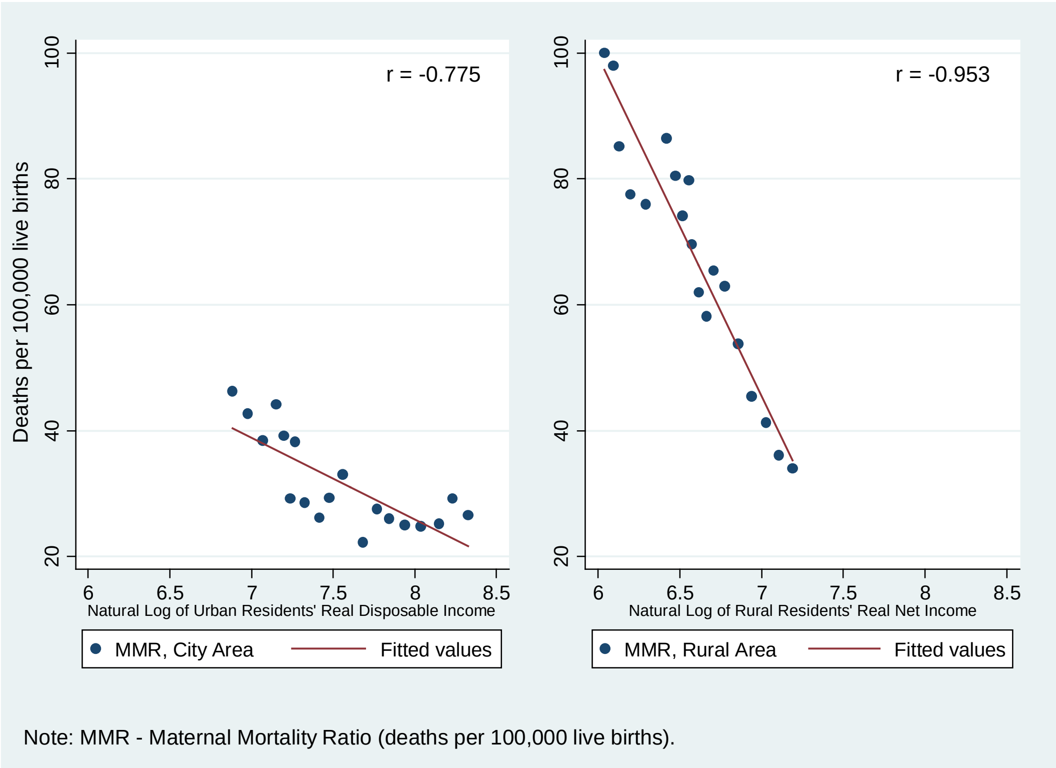


Figure 5.1. The correlation between maternal mortality ratio and real personal income in China, 1991-2009.

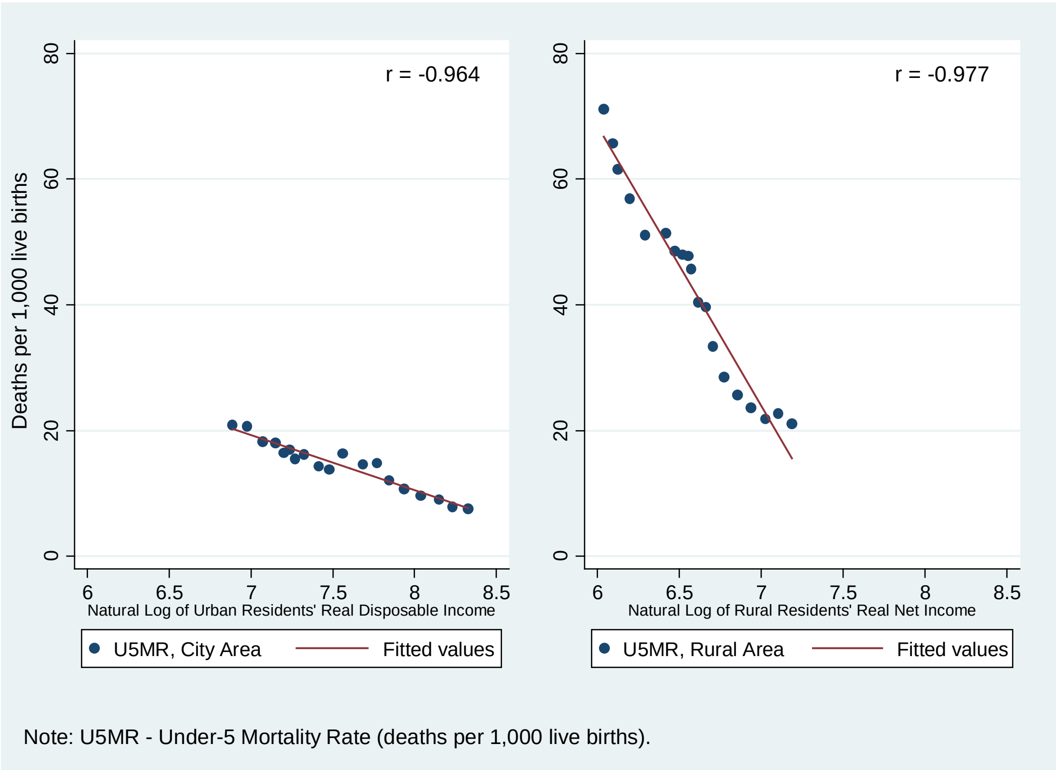


Figure 5.2. The correlation between under-five mortality rate and real personal income in China, 1991-2009.

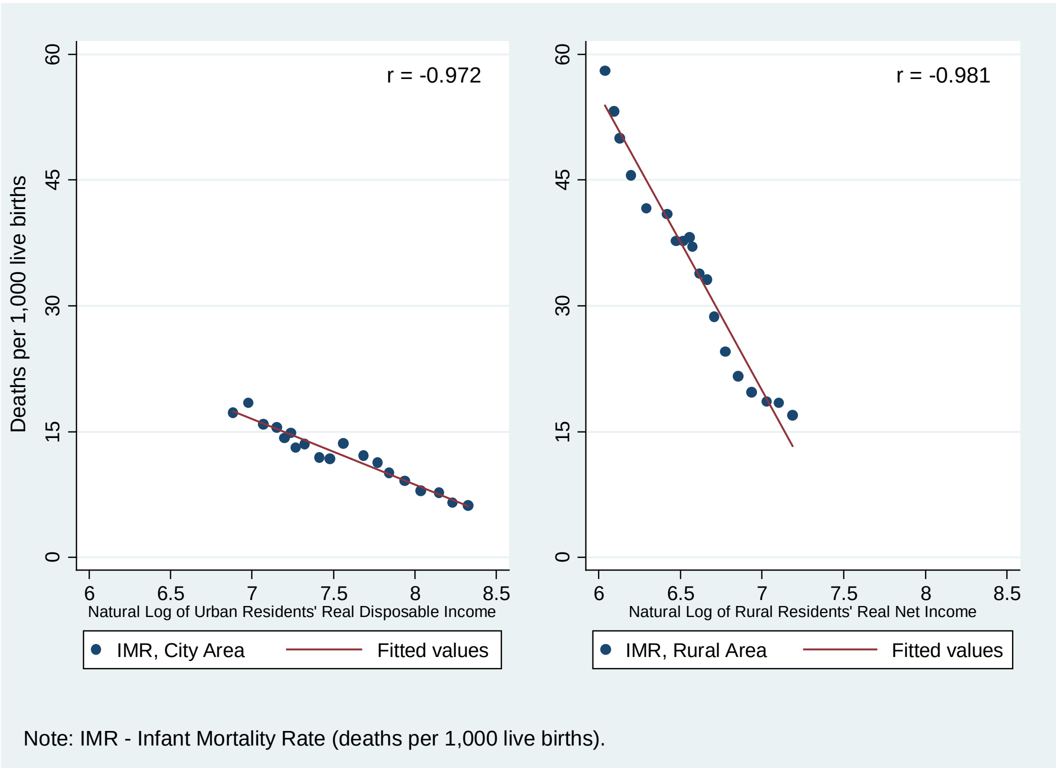


Figure 5.3. The correlation between infant mortality rate and real personal income in China, 1991-2009.

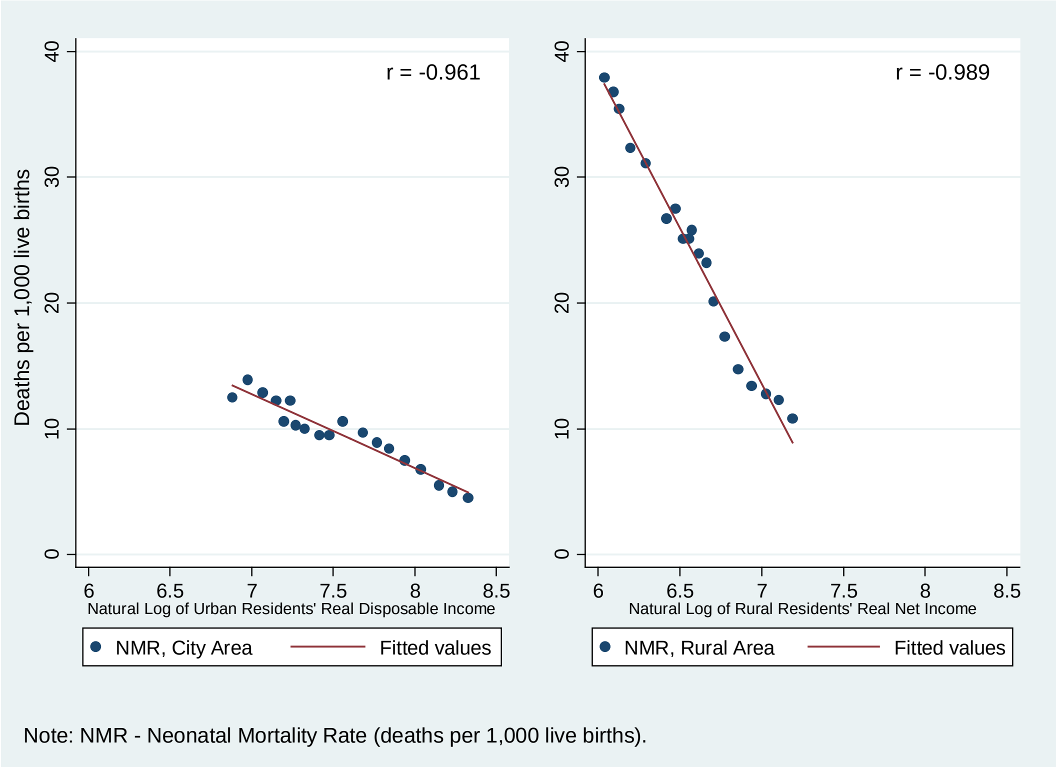


Figure 5.4. The correlation between neonatal mortality rate and real personal income in China, 1991-2009.

5.3.1.3. Bivariate correlation between economic fluctuation and health outcomes

The correlation between the four mortality indicators and economic fluctuations (proxied by the changes of the natural log of real personal income in urban and rural area) are presented in Figures 5.5-5.8. As is shown, there is a negative relationship between real personal income growth and mortality, while the correlation is slightly stronger in urban China.

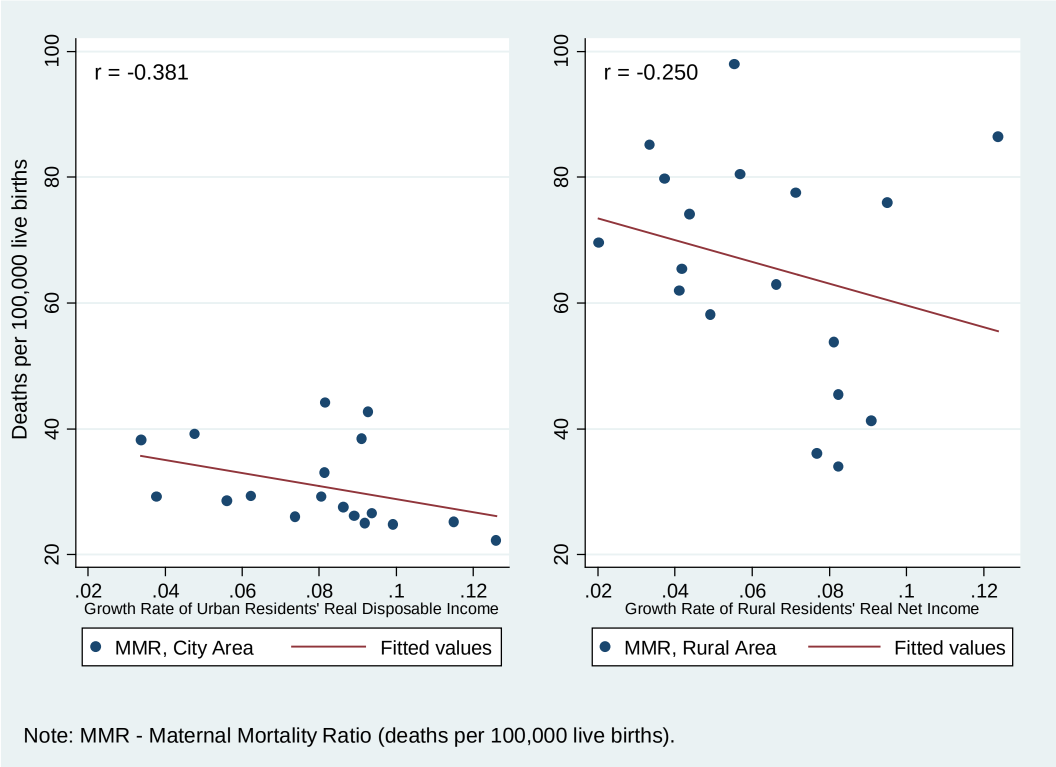


Figure 5.5. The correlation between maternal mortality ratio and real personal income fluctuation in China, 1991-2009.

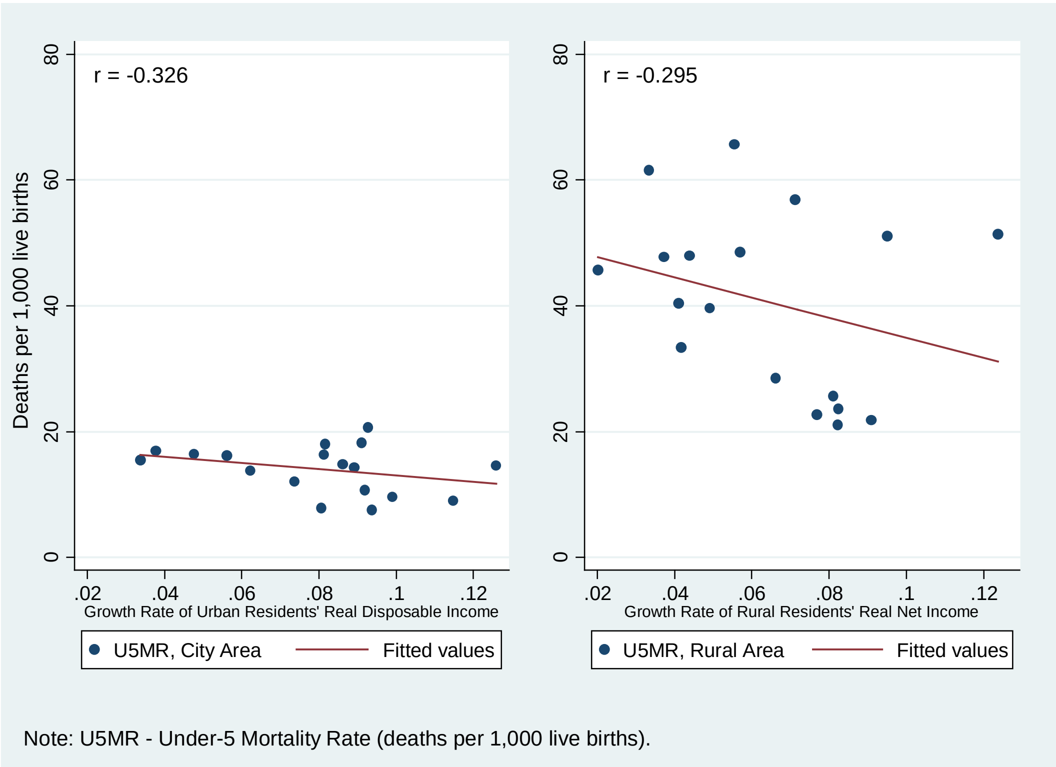


Figure 5.6. The correlation between under-five mortality rate and real personal income fluctuation in China, 1991-2009.

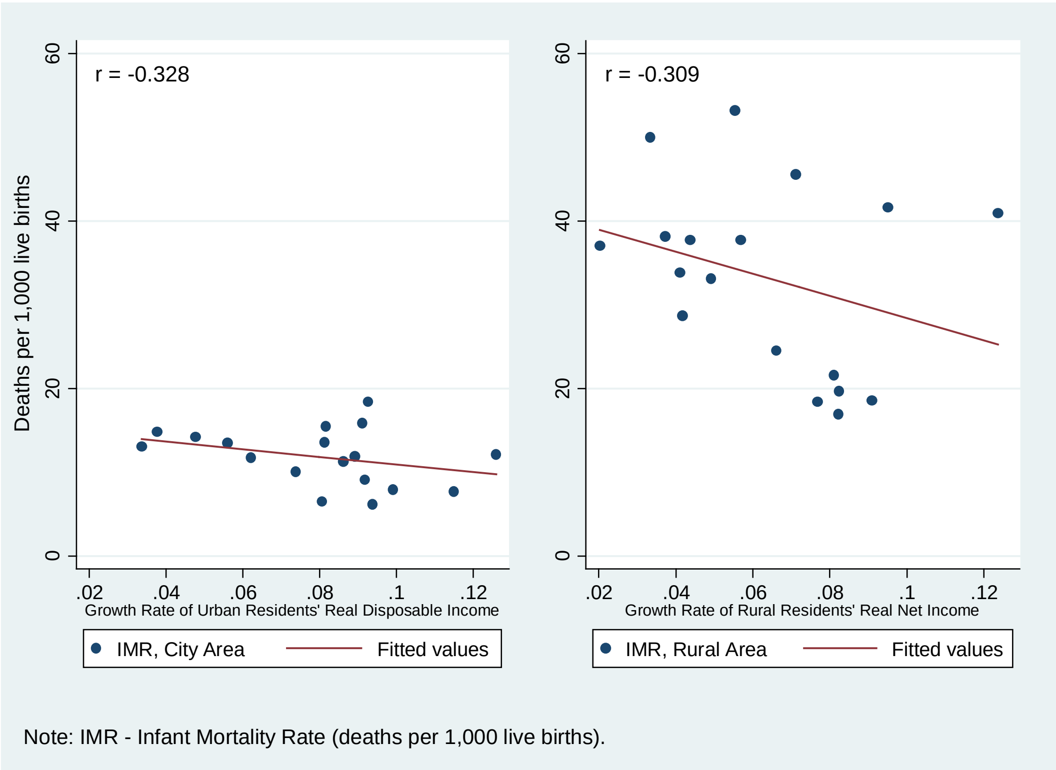


Figure 5.7. The correlation between infant mortality rate and real personal income fluctuation in China, 1991-2009.

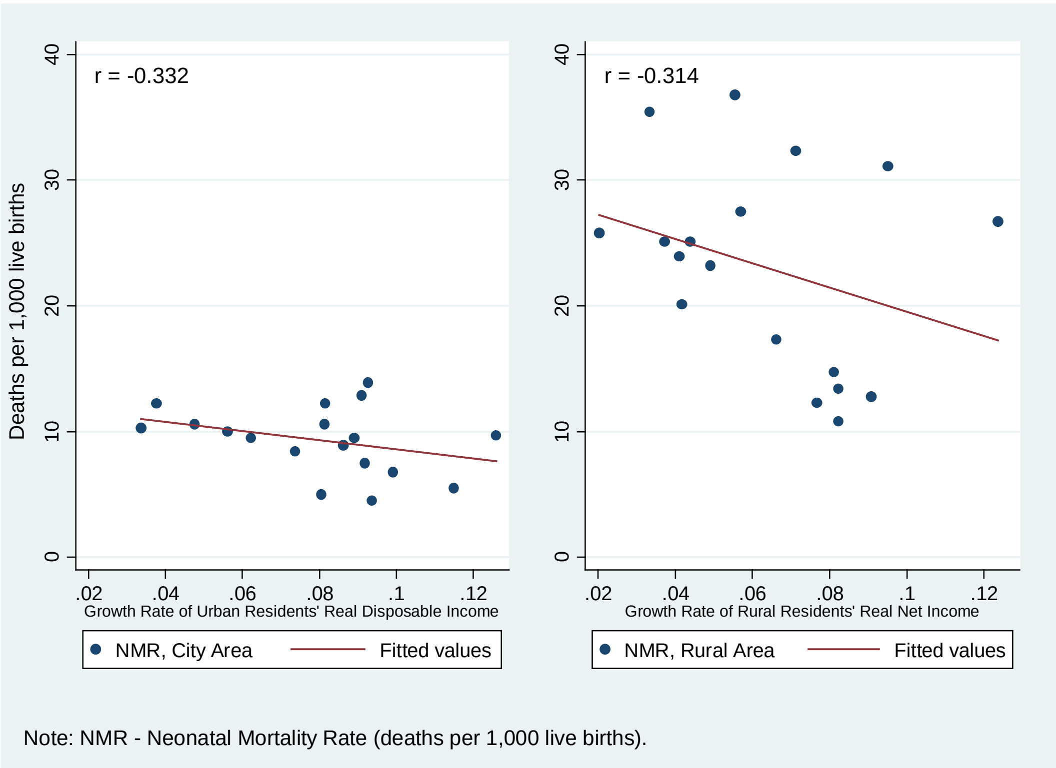


Figure 5.8. The correlation between neonatal mortality rate and real personal income fluctuation in China, 1991-2009.

5.3.2. Correlation Analysis from Provincial-level Panel Data

5.3.2.1. Intra-provincial disparities in health outcomes

From the descriptive analyses reported in the last section, rural/urban disparities can be clearly observed. In this section, by utilising the provincial level panel data, further intra-provincial disparities are found in the heath outcome indicators.

Rather than plotting time series for all provinces, five provinces have been selected to depict the trends for the health outcomes measures. In addition to Beijing (which is the capital city) and Tibet (which is often regarded as an outlier in empirical analysis), the other three provinces which had the highest (Xinjiang), middle (Jiangxi) and lowest (Shanghai) MMR in 2002 are selected.

In the two maternal and child mortality indicators plotted in Figures 5.9 and 5.10, it can be seen that there is a downward trend for intra-provincial health inequality in MMR. It can also be observed that the gap between the best and worst performing provinces for PNM has increased from 2002 to 2008.

As for the other two child nutrition related indicators presented in Figures 5.11 and 5.12, the observed trend is not so optimistic. For example, although the intra-provincial health inequality has slightly decreased, the time-series for low birth weight variables actually show a worsening outcome during this eight-year period.

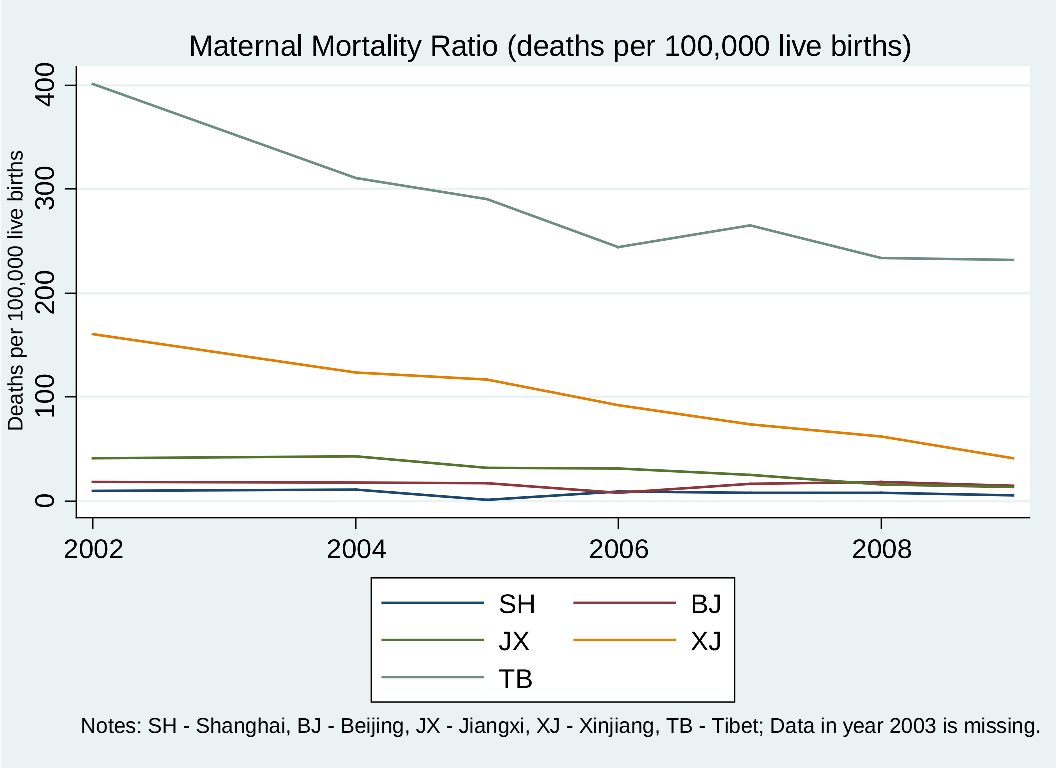


Figure 5.9. The maternal mortality rate in selected provinces, 2002-2009.

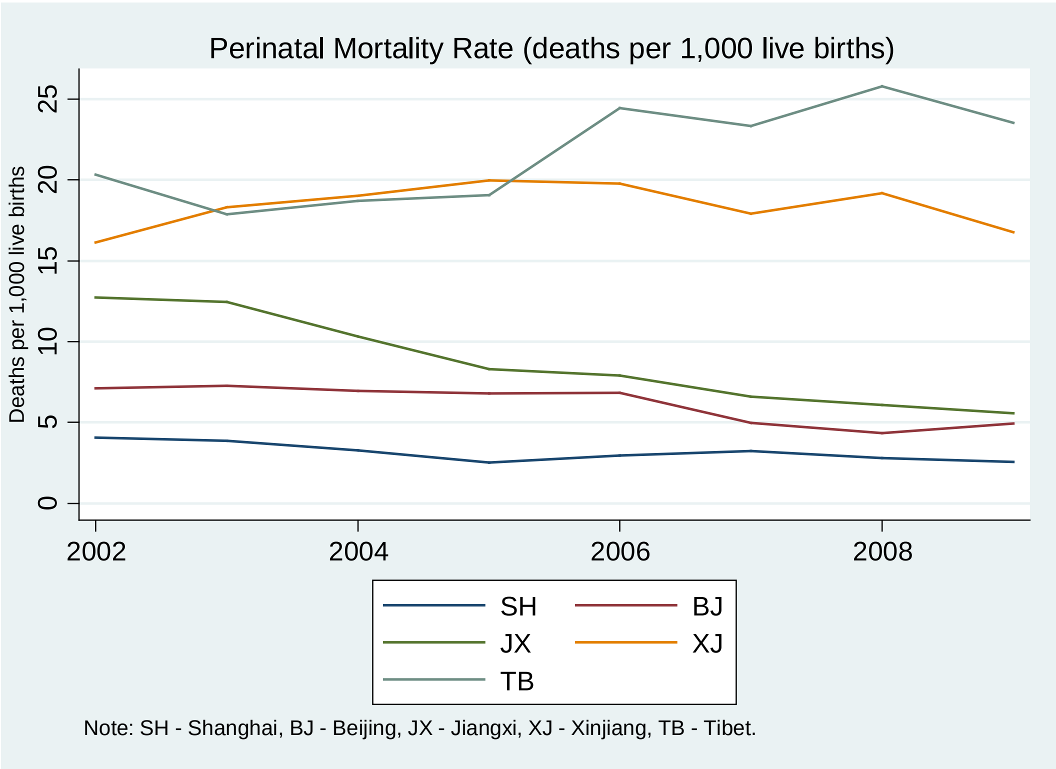


Figure 5.10. The perinatal mortality rate in selected provinces, 2002-2009.

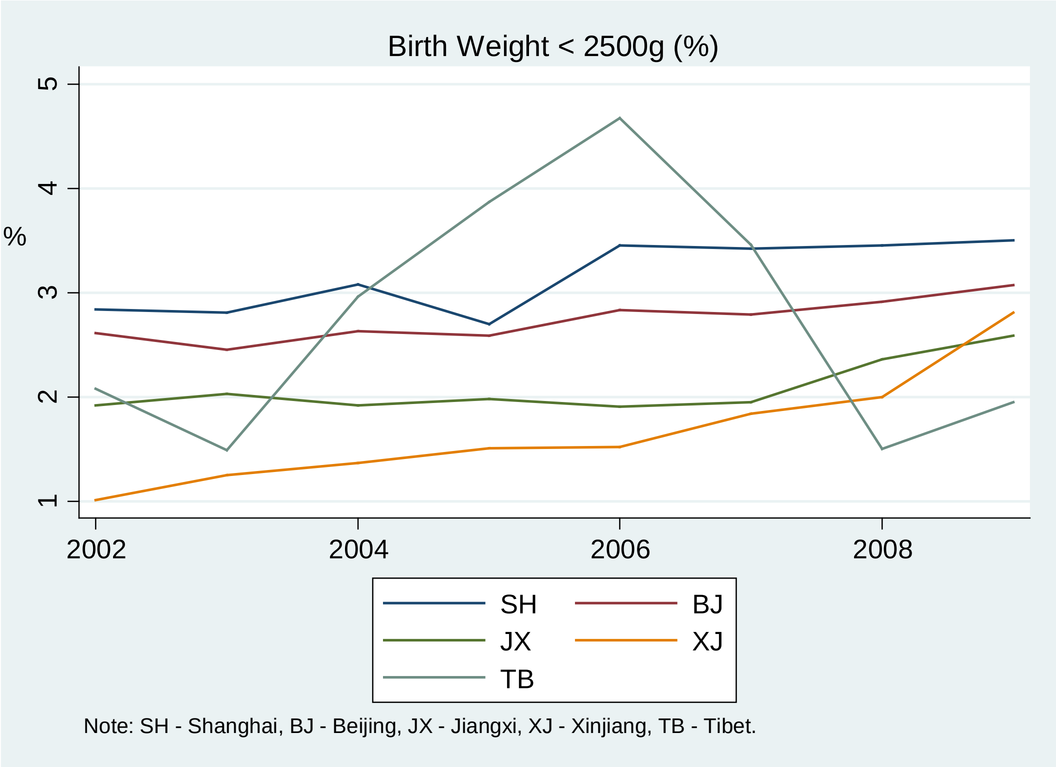


Figure 5.11. The percentage of low birthweight newborns in selected provinces, 2002-2009.

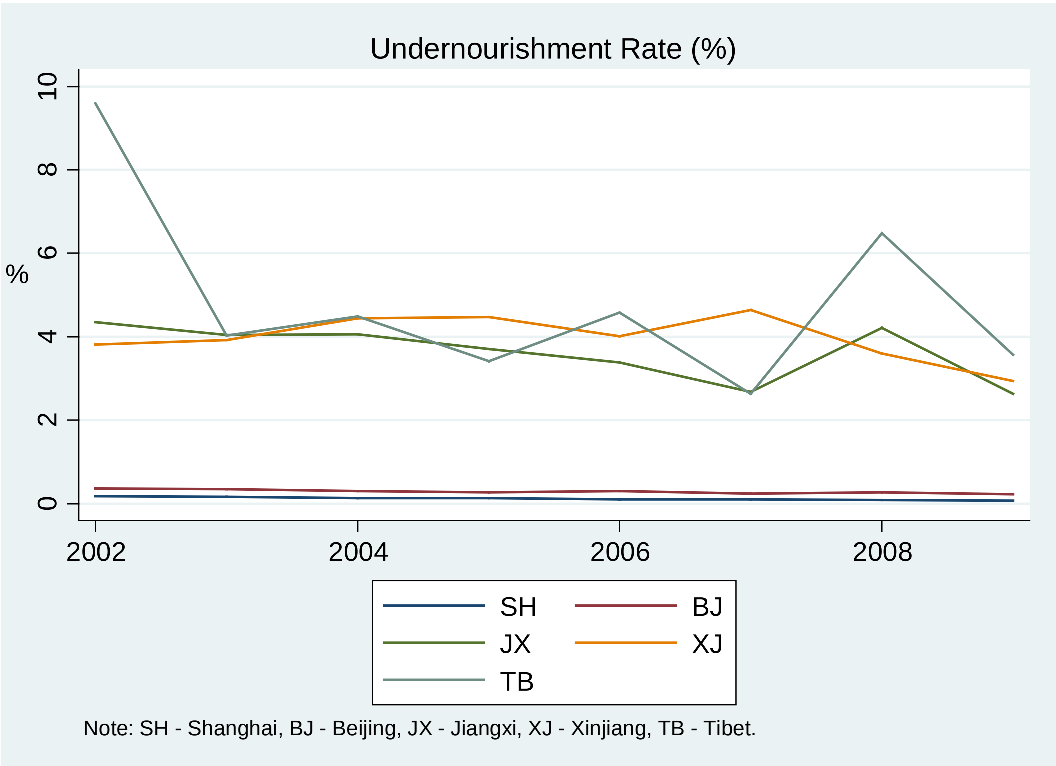


Figure 5.12. The undernourishment rates among children under 5 years age in selected provinces, 2002-2009.

The mortality rate and incidence rate of TB are shown in Figures 5.13 and 5.14. Two patterns can be observed. Firstly, although there are several fluctuations among the series, there is an increasing trend for both mortality and incidence rates in three provinces as plotted in the figure. Secondly, intra-provincial health inequality relating to TB prevalence has slightly increased.

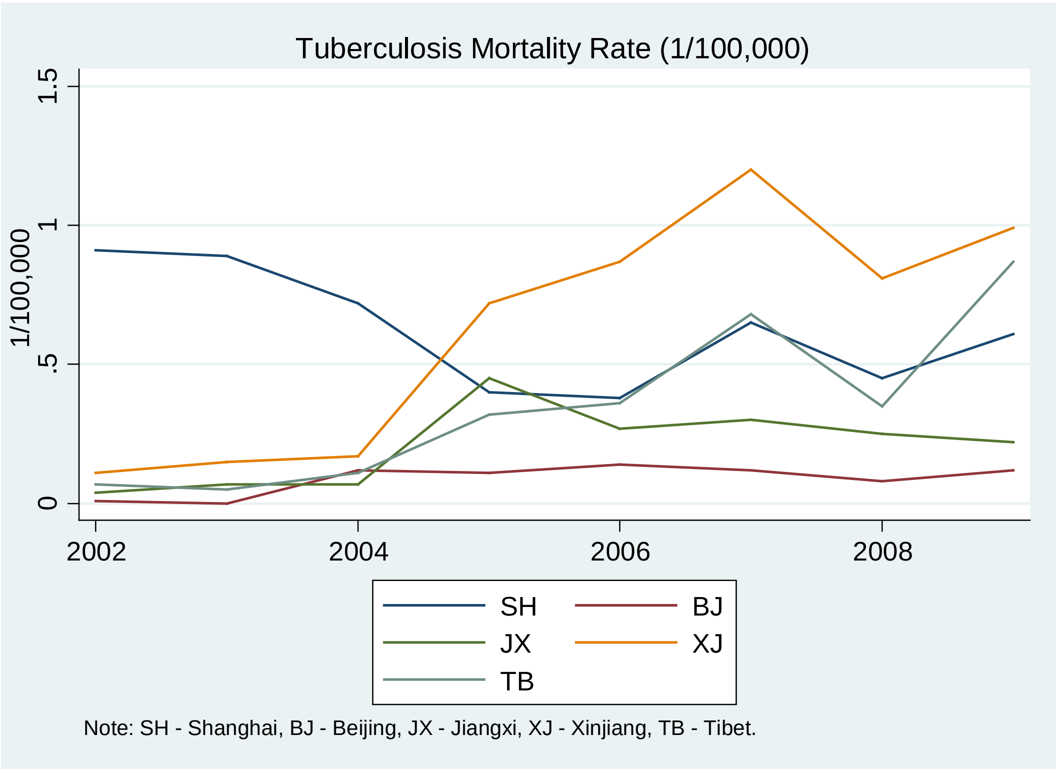


Figure 5.13. The TB mortality rate in selected provinces, 2002-2009.

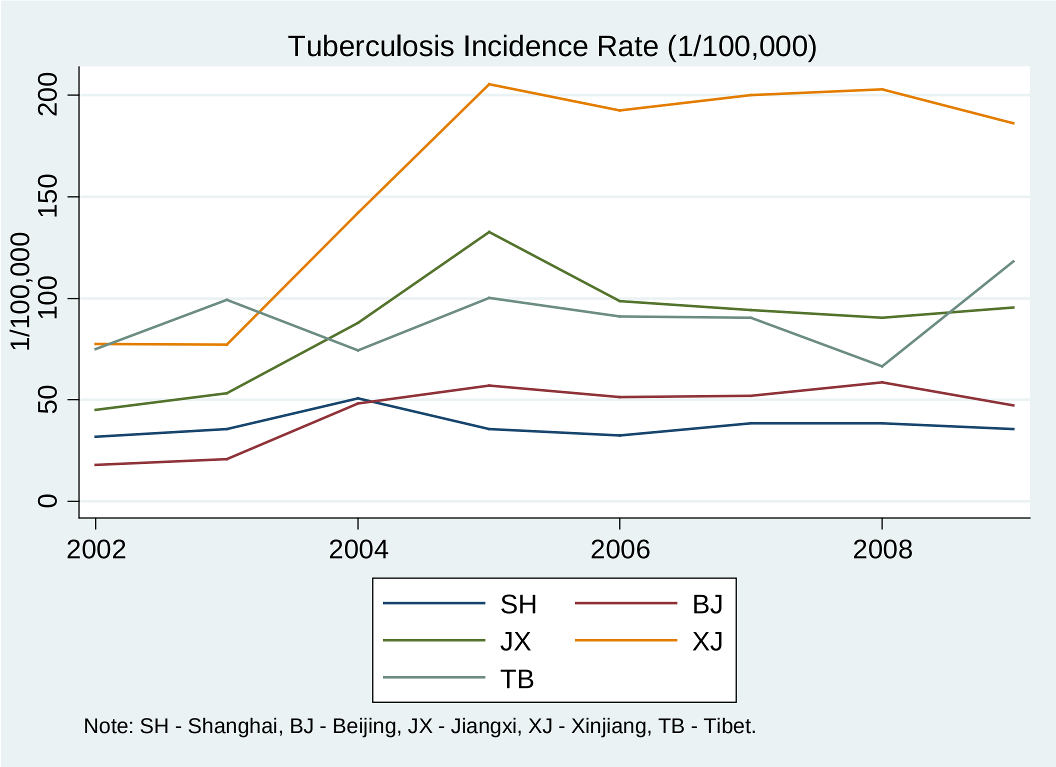


Figure 5.14. The TB incidence rate in selected provinces, 2002-2009.

5.3.2.2. Bivariate correlation between economic development, economic fluctuations and health outcomes

Before reporting the regression results, the correlation between four maternal and child health outcome indicators, two infectious diseases indicators (dependent variable) and the economic development/fluctuation indicators are plotted in Figures 5.15 to 5.18, using pooled provincial-level time series data.

Briefly, the level of economic development is associated with a lower maternal and child mortality, malnutrition level and the incidence of TB. However, not all health outcome indicators improve along with the increased economic output levels: low birth weight and TB mortality rate clearly fit this group.

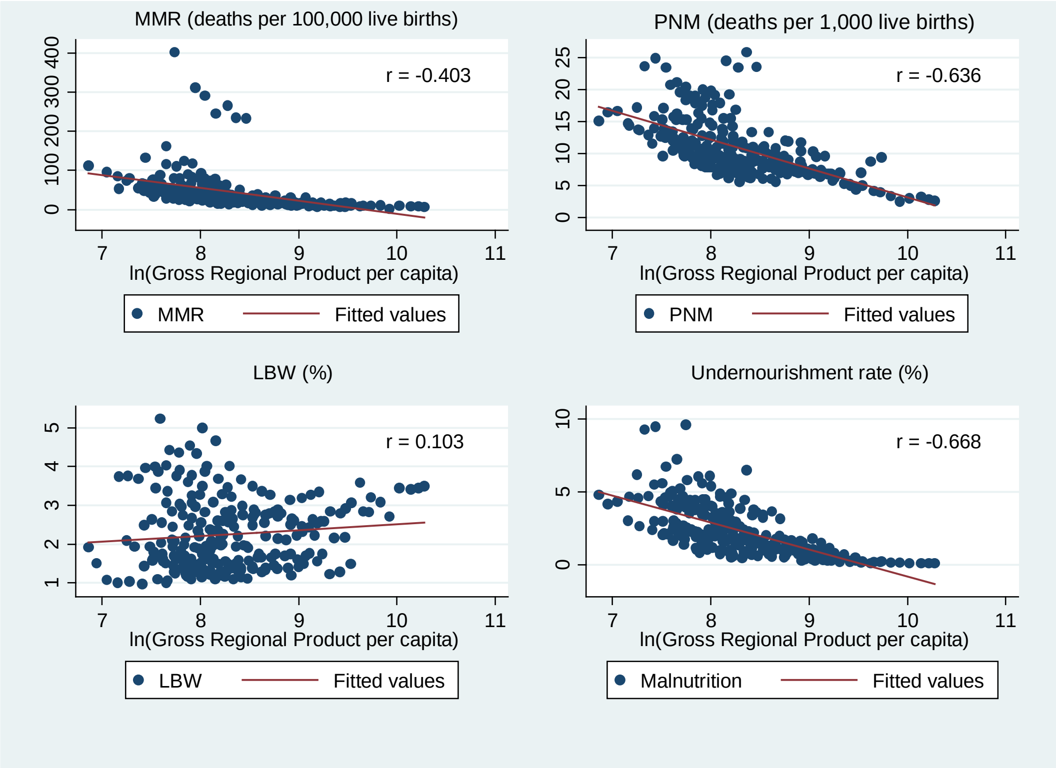


Figure 5.15. The correlation between maternal and child health outcome indicators and economic development.

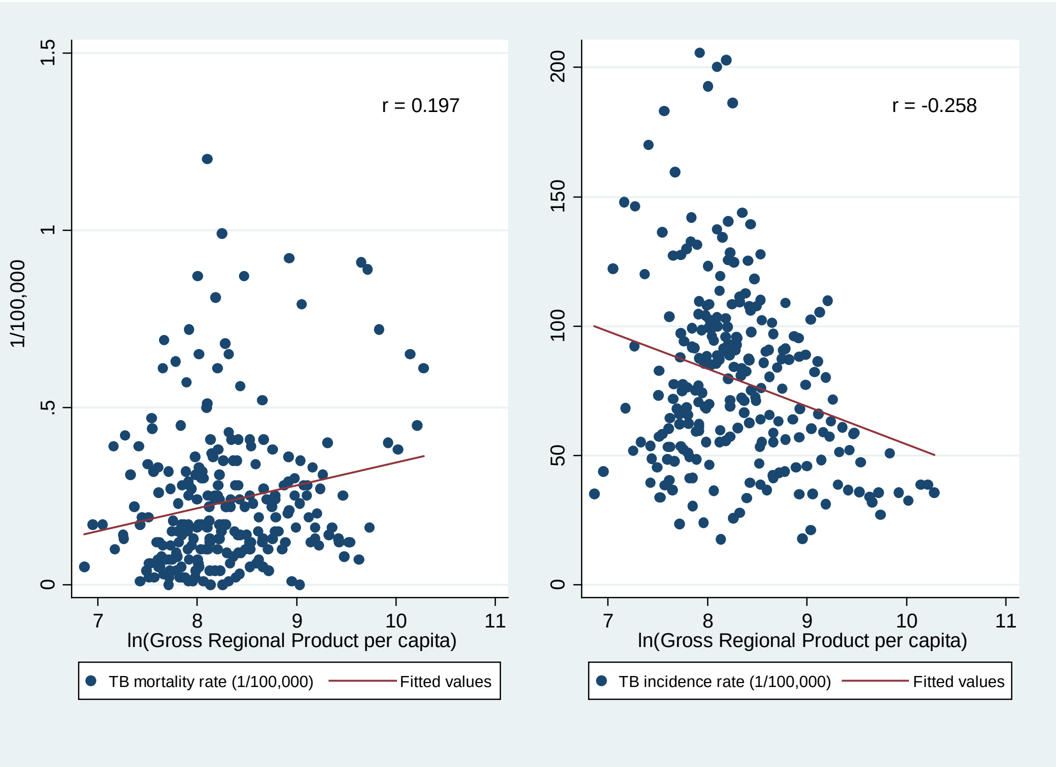


Figure 5.16. The correlation between infectious disease indicators and economic development.

The relationship between health outcomes and economic fluctuations are quite different between maternal and child health indicators, and infectious diseases indicators. Positive economic shocks (i.e. higher real GDP per capita growth) are associated with better maternal and child health, while on the contrary, worse infectious disease is prevalent.

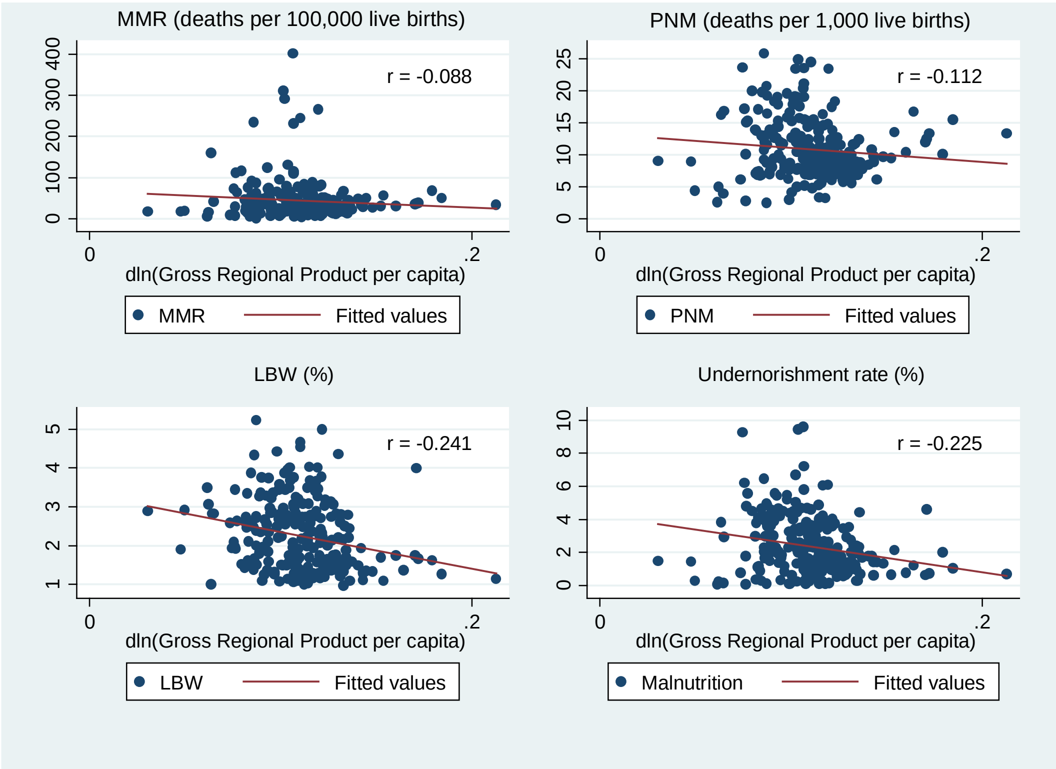


Figure 5.17. The correlation between maternal and child health outcome indicators and economic fluctuations.

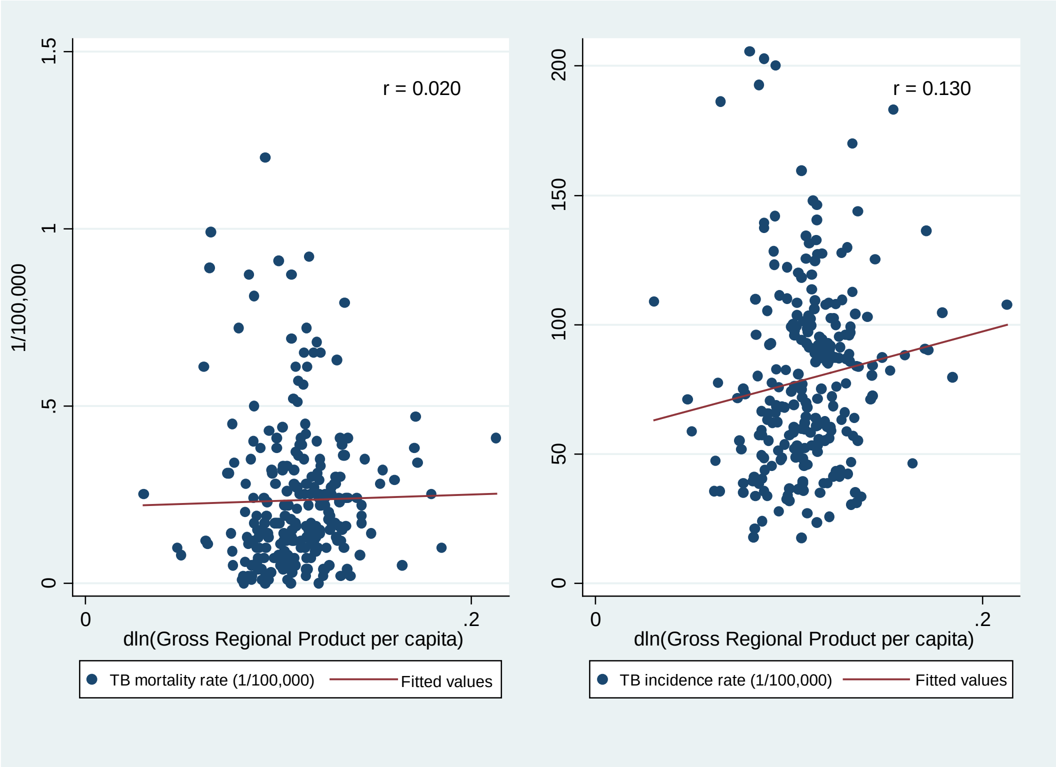


Figure 5.18. The correlation between infectious disease indicators and economic fluctuations.

5.3.3. Regression Analysis from Provincial-level Panel Data

Fixed effect estimates are reported in the following tables. For each health outcome indicator, there are two groups (tables) of results reported based on Equations 5.1 and 5.2 separately. The dependent variables are the original form of health outcome indicators (i.e. levels).

5.3.3.1. Economic development, health investment and maternal health outcomes measurements

The relationship between economic development, health investment and maternal mortality ratio is reported in Table 5.3. The results based on a whole China sample are reported in the first two columns. The real GRP per capita is used as the proxy for economic development in the whole China sample. It is expected that economic development will improve health status (as reflected by a significant negative coefficient). The estimated coefficient is negative as expected, but it is insignificant. In the second column, an interaction term between ln(GRP per capita) and a dummy variable indicating the sample comes from inland China is included. The estimated coefficient is significantly negative, suggesting that economic development is significantly associated with an improved maternal health status in the inland China sample. Since a right-hand side semi-log function form is adopted, holding other covariates constant, a 1% increase in real GRP per capita reduced the MMR by around 23.2 per 100,000 live births in inland China.

Health investment and health service access are also key variables of interest in this study. Except for the percentage of government health expenditure as a share of total local government health expenditure, which is included as a general measure of health investment in each province, a more direct MMR related health services utilisation variable, modern birth attendance rate is specifically controlled for, and it is hypothesised that both coefficients would be negative. As can be found in the first two columns, an increase in the number of modern birth attendances significantly reduced the MMR. General government health investment is insignificant. Other control variables included in the MMR equation are all insignificant.

Since disaggregated MMR data in urban and rural areas are available, a sub-sample analysis is performed to seek robust results. It should be noted that in the sub-sample analysis only urban and rural specific control variables are included. For example, since the disaggregated GRP variable is unavailable, the personal income of urban and of rural residents are included separately for urban and rural samples. Certain control variables (i.e. general government investment, age structure and education level) are not included due to data availability. This may induce omitted variable bias. However, as can be found in the whole China sample, those variables are all insignificant so it is argued that the bias is likely to be small.

The sub-sample results show that firstly, economic development is significantly related to the MMR: in both the urban and rural China sample, increased real personal income is significantly associated with a decreased MMR, the magnitude of the coefficient is stronger in rural China compared to urban China. The interaction terms between real personal income and the inland China dummy confirms what has been found in the whole China sample; that the relationship between economic development and MMR is stronger in inland China. The health services access variable, modern birth attendance rate, is robustly significantly negative, telling the same story as that which has been identified based on the whole China sample.

Table 5.3. Fixed effects estimates of factors on maternal mortality, 2004-2009.

	whole China		urban China		rural China	
	(1)	(2)	(3)	(4)	(5)	(6)
ln(ECO)	-15.726	-10.651	-17.228***	-9.729**	-44.105***	-26.169***
	[11.446]	[10.847]	[5.390]	[4.658]	[7.948]	[7.952]
ln(ECO)*Inland		-23.238**		-13.798*		-32.769**
		[10.155]		[7.154]		[15.113]
Modern Birth Attendance	-4.328***	-3.935***	-4.863**	-4.721**	-4.299***	-3.951***
	[0.486]	[0.572]	[1.941]	[2.065]	[0.649]	[0.645]
MED_GOV	-12.629	117.620				
	[114.593]	[136.831]				
POP014	22.705	-13.264				
	[69.904]	[72.626]				
POP65P	-220.065	-147.828				
	[163.517]	[153.340]				
URB	-53.557	-48.241				
	[56.312]	[58.920]				
ASY	-0.774	-2.946				
	[4.746]	[4.094]				
Obs.	186	186	186	186	185	185
Provinces	31	31	31	31	31	31

Notes: Bootstrap standard errors in brackets (the bootstrap is based on 1000 redraws and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1. ECO refers to real GRP per capita for the whole China sample, per capita real annual disposable/net income of urban/rural household for the urban/rural China sample.

5.3.3.2. Economic development, health investment and child health outcomes measurements

The FE estimates for low birth weight, perinatal mortality rate and undernourishment rates among children under 5 years of age are reported in Table 5.4. Among three health outcome indicators, economic development is found to be significantly associated with PNM and undernourishment rates. A 1% increase in real GRP per capita was associated with a decreased PNM of around 2.4 deaths per 1,000 live births, or a decreased undernourishment rate of around 1.6% holding other variables constant. For malnutrition, there is also some evidence that the increased economic levels are associated with decreased malnutrition in inland China (as seen by the interaction term).

As with the MMR equation, except for general government health investment measures, two health service access variables (antenatal examination rate and neonatal supervision rate) are included to investigate the relationship between health services access and health outcomes (LBW and PNM). The estimated coefficient is consistently negative, but insignificant. Regarding general government health spending, there is some evidence suggesting increased government health spending is significantly associated with a decreased PNM.

Most other covariates are insignificant, except for the urbanisation rate (in the PNM equation) and education level (in the LBW equation). The significant negative coefficient of the urbanisation rate in the PNM equation is as expected. The significant positive coefficient of education levels in the LBW equation is not expected. A higher education level would be expected to lead to a higher income, better health related knowledge, healthier behaviour and thus a better health outcome. However, as has been discussed in section 5.3.2.1, low birth weight exhibits a different trend in China than the other three maternal and child health outcome indicators: despite improved living standards, low birth weights have increased slightly. The unexpected sign of education levels on child health outcomes has also been documented elsewhere. Using macro data from Taiwan, Lin (2006) has also reported a similar result based on a different proxy for education, the percentage of female population graduated from college. As discussed in Lin (2006), the increased education level and the increased labor force participation rate may delay the marrying age and the age of reproduction. Increased maternity age is associated with a higher risk for both pregnant women and babies. Gennetian *et al*. (2010) and Morrill (2011) also found adverse effect of maternal employment on children's health status using the survey data from United States. More evidence using survey data from China would be useful to investigate this further.

Table 5.4. Fixed effects estimates of factors on child health outcomes, 2002-2009.

	LBW		PNM		MAL	
	(1)	(2)	(3)	(4)	(5)	(6)
ln(GRP per capita)	0.499	0.756	-2.360*	-1.326	-1.608*	-1.038
	[0.503]	[0.582]	[1.377]	[1.170]	[0.822]	[0.755]
ln(GRP per capita)*Inland		-0.522		-2.112		-1.165**
		[0.450]		[1.355]		[0.496]
Antenatal Examination	0.005	0.005				
	[0.019]	[0.019]				
Neonatal Supervision			-0.040	-0.040		
			[0.053]	[0.052]		
MED_GOV	-2.510	0.234	-47.385**	-36.367*	-12.829	-6.751
	[6.343]	[5.524]	[20.909]	[20.822]	[8.081]	[8.979]
POP014	7.066	6.592	-19.417	-21.318	-3.955	-5.004
	[4.882]	[4.511]	[16.940]	[17.220]	[4.848]	[5.109]
POP65P	-2.666	-0.212	-14.111	-4.087	-2.336	3.193
	[6.871]	[7.906]	[21.688]	[23.372]	[10.239]	[10.556]
URB	-1.477	-2.130	-18.802**	-21.435**	2.361	0.908
	[2.647]	[2.784]	[9.438]	[8.732]	[3.743]	[3.357]
ASY	0.485**	0.459**	-0.391	-0.497	-0.099	-0.157
	[0.200]	[0.192]	[0.830]	[0.851]	[0.424]	[0.420]
Y2003	-0.080	-0.095	0.428	0.369	0.094	0.062
	[0.085]	[0.083]	[0.318]	[0.299]	[0.185]	[0.186]
Observations	248	248	248	248	248	248
Provinces	31	31	31	31	31	31

Notes: Bootstrap standard errors in brackets (the bootstrap is based on 1000 redraws and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

5.3.3.3. Economic development, health investment and infectious disease measurement

The results regarding TB mortality rates and incidence rates are reported in Table 5.5. Consistent with the correlation analysis reported in Figure 5.5.4, the coefficient on economic level is positive, although insignificant. As for the relationship with TB incidence rates, there is some evidence suggesting the association is significant in inland China. Holding other variables constant, a 1% increase in real GRP per capita was associated with an increased incidence rate at around 33.5 per 100,000 people in inland China. The association with government health investment is similar to the maternal and child health outcomes, where increased government health expenditure is associated with a decreased TB mortality and incidence rate, the relationship is significant in the TB incidence equation.

Other significant covariates suggest that increased education level is associated with decreased mortality, and the urbanisation rate is associated with an increased TB incidence rate. The year dummy of 2003 suggests that during the SARS shock, the TB incidence rate is significantly lower.

Table 5.5. Fixed effects estimates of factors on TB, 2002-2009.

	TB mortality		TB incidence	
	(1)	(2)	(3)	(4)
ln(GRP per capita)	0.159	0.086	39.628	23.227
	[0.222]	[0.263]	[33.098]	[29.333]
ln(GRP per capita)*Inland		0.150		33.505**
		[0.141]		[14.353]
MED_GOV	-0.039	-0.819	-778.375***	-953.126***
	[2.258]	[1.728]	[245.511]	[247.784]
POP014	-0.064	0.071	-83.400	-53.242
	[1.677]	[1.689]	[263.072]	[277.778]
POP65P	3.579	2.870	367.788	208.806
	[2.568]	[2.623]	[296.010]	[330.255]
URB	0.663	0.850	210.082	251.841*
	[1.056]	[1.059]	[152.046]	[136.571]
ASY	-0.152**	-0.145**	-14.619	-12.951
	[0.072]	[0.068]	[9.729]	[9.148]
Y2003	-0.010	-0.006	-7.367***	-6.433***
	[0.020]	[0.020]	[2.347]	[2.418]
Observations	248	248	248	248
Provinces	31	31	31	31

Notes: Bootstrap standard errors in brackets (the bootstrap is based on 1000 redraws and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

5.3.3.4. Robust analysis for the relationship between economic development and health outcomes

The key results using an instrumental variable approach are summarised in Table 5.6. In all the columns the Kleibergen-Paap *rk* Wald *F* statistics are reported and, using the same standard adopted in Chapter 4, they all suggest the instruments are not weak. The first stage result of Column (1), Panel A, Table 5.6 is used as an example to show that the fixed assets investment is a valid instrument. The Kleibergen-Paap *rk* Wald *F* statistic is already reported in the table and thus not replicated here. The first stage estimated result for ln(GRP per capita) equation is:

$$\ln(GRP \;\; per \;\; capita_{it}) = \underset{[0.034]^{***}}{0.384} \times \ln(INV \;\; per \;\; capita_{it}) + z_{1it}$$

where the HAC robust standard errors are reported in brackets, the coefficient on fixed capital investment is positive and significant at the 1% level, the detailed results of other variables included in the z_1 vector (Modern Birth Attendance, MED_GOV, POP014, POP65P, URB, ASY) are not reported for simplicity. As anticipated by theory, fixed capital investment has a significant and positive effect on GDP.

Whether the income variable is endogenous in each health outcome equation is tested by a Hausman test. The endogeneity test statistics are reported in the Table 5.6. Conditional on the validity of the instrument, it can be concluded that in most health outcomes equations, the null hypothesis that the economic development variable is exogenous cannot be rejected at the 10% level. The only three exceptions are the equations reported in Columns (2) and (4) in Panel A, where MMR or MMR in urban areas are the dependent variables and the equation reported in Column (2) in Panel C where the TB mortality rate is the dependent variable. This result is consistent with the discussion in the methodology section that when child health outcomes are the dependent variables, reverse causality of the income variable is less of an issue. This also supports the results reported by Jin and Sun (2011) that income can be regarded as exogenous when IMR is studied.

Comparing the IV/GMM estimates with FE estimates, it is not surprising that in most cases results are comparable (since economic development could be regarded as exogenous in most equations). When economic development is exogenous in the equation, the FE estimates are more efficient, thus the conclusion is analogous to the discussion above. There is only one change when income variables are treated as endogenous and the statistical significance of a variable changes: the interaction term between income and inland China in urban China is now insignificant for MMR indicators (Column (4), Panel A).

Table 5.6. IV/GMM estimates of factors on health outcomes.

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: MMR						
	whole China		urban China		rural China	
ln(ECO)	-20.111*	-11.951	-18.820**	-13.906***	-48.891***	-30.286***
	[11.642]	[10.518]	[8.833]	[4.786]	[6.965]	[6.555]
ln(ECO)*Inland		-19.536***		-7.821		-29.483***
		[6.148]		[11.595]		[11.101]
Kleibergen-Paap *rk* Wald *F* statistic	127.741	67.549	763.791	134.337	879.659	110.172
Endogeneity test statistics	0.208	4.973*	0.293	5.474*	2.391	0.857
Observations	186	186	186	186	185	185
Panel B: Child health outcomes						
	LBW		PNM		MAL	
ln(GRP per capita)	0.253	0.528	-2.016	-0.719	-1.572*	-0.956
	[0.571]	[0.629]	[1.566]	[1.406]	[0.843]	[0.739]
ln(GRP per capita)*Inland		-0.455		-2.143**		-1.018***
		[0.346]		[0.852]		[0.349]
Kleibergen-Paap *rk* Wald *F* statistic	239.904	116.466	235.119	113.841	241.017	116.869
Endogeneity test statistics	0.590	1.680	0.085	0.891	0.006	3.826
Observations	248	248	248	248	248	248
Panel C: TB						
	TB mortality		TB incidence			
ln(GRP per capita)	0.234	0.177	20.401	2.549		
	[0.193]	[0.229]	[21.382]	[20.270]		
ln(GRP per capita)*Inland		0.095		29.529***		
		[0.092]		[9.218]		
Kleibergen-Paap *rk* Wald *F* statistic	241.017	116.869	241.017	116.869		
Endogeneity test statistics	0.804	5.898*	2.033	3.133		
Observations	248	248	248	248		
Provinces	31	31	31	31	31	31

Notes: Heteroskedasticity-autocorrelation consistant (HAC) robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. ECO refers to real GRP per capita for the whole China sample, per capita real annual disposable/net income of urban/rural household for the urban/rural China sample. Only the coefficients of endogenous variables are reported here, full control variables please refer to Tables 5.4-5.5. The Kleibergen-Paap *rk* Wald *F* statistic suggest that the instruments are not weak.

5.3.3.5. Economic fluctuations and health outcomes measurements

The relationship between economic fluctuations and health outcomes is reported in Tables 5.7 to 5.9. As mentioned in the methodology section, the first difference of ln(GRP per capita) is used to proxy for economic fluctuations. An interaction term between economic fluctuations and the inland region dummy is also considered. Other variables are the same as included in the health outcomes level equations. In this model setting, the pure effect of short-run economic fluctuations is estimated conditional on the level of living standards.

In the following three tables, the economic fluctuations variable is found to be insignificant in most of the models estimated. Two exceptions are the PNM and TB incidence rates. Assuming the association is symmetric, there is some evidence suggesting that a positive economic shock is significantly associated with a reduced PNM after controlling the level of living standards, health investment and other covariates (Table 5.8, Columns (3) & (4)). The relationship between economic fluctuations and TB incidence rates is contrary to what has been identified for the PNM, where a positive economic shock is associated with an increased TB incidence (Table 5.9, Columns (3) & (4)). One possible explanation is that a positive economic shock may increase daily activities, population density and further increase the likelihood of transmission. A similar phenomenon can be observed in Great Britain and Japan during rapid industrialisation and urbanisation processes (Lönnroth *et al*., 2010). In this analysis, the urbanisation rate has also been included directly into the model and the estimated coefficient is positive in all TB incidence equations, while significant in some occasions (i.e. Column (4) in Tables 5.5 and 5.8). As China is undergoing both economic growth and urbanisation, their different effects on TB incidence cannot be distinguished clearly in the analysis.

Table 5.7. Fixed effects estimates of economic fluctuations on maternal mortality, 2004-2009.

	whole China		urban China		rural China	
	(1)	(2)	(3)	(4)	(5)	(6)
Δln(ECO)	-0.648	-9.792	142.047	-16.282	-8.329	-1.808
	[33.841]	[35.588]	[112.991]	[37.501]	[30.628]	[66.314]
Δln(ECO)*Inland		43.671		235.352		-20.512
		[71.803]		[163.531]		[74.402]
ln(ECO)	-15.731	-11.494	-24.233***	-11.736***	-43.888***	-26.122***
	[11.572]	[11.397]	[5.800]	[3.683]	[8.040]	[6.765]
ln(ECO)*Inland		-22.884**		-23.584**		-32.406**
		[10.428]		[11.279]		[14.454]
Modern Birth Attendance	-4.328***	-3.941***	-3.803***	-3.215***	-4.294***	-3.941***
	[0.485]	[0.569]	[1.069]	[0.926]	[0.650]	[0.651]
MED_GOV	-12.453	114.651				
	[114.636]	[135.292]				
POP014	22.773	-22.779				
	[70.268]	[76.401]				
POP65P	-219.680	-151.762				
	[167.312]	[157.568]				
URB	-53.717	-43.641				
	[57.407]	[62.161]				
ASY	-0.779	-3.063				
	[4.716]	[4.116]				
Obs.	186	186	186	186	185	185
Provinces	31	31	31	31	31	31

Notes: Bootstrap standard errors in brackets (the bootstrap is based on 1000 redraws and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1. ECO refers to real GRP per capita for the whole China sample, per capita real annual disposable/net income of urban/rural household for the urban/rural China sample.

Table 5.8. Fixed effects estimates of economic fluctuations on three child health outcomes, 2002-2009.

	LBW		PNM		MAL	
	(1)	(2)	(3)	(4)	(5)	(6)
Δln(GRP per capita)	-2.397	-2.107	-8.790*	-5.904	-3.163	-0.063
	[2.143]	[2.526]	[5.207]	[7.318]	[2.590]	[2.151]
Δln(GRP per capita)*Inland		0.560		-0.881		-3.213
		[3.536]		[14.872]		[4.808]
ln(GRP per capita)	0.548	0.767	-2.178	-1.257	-1.544*	-0.993
	[0.491]	[0.581]	[1.394]	[1.221]	[0.819]	[0.778]
ln(GRP per capita)*Inland		-0.488		-1.959		-1.088**
		[0.454]		[1.474]		[0.527]
Antenatal Examination	0.005	0.005				
	[0.019]	[0.018]				
Neonatal Supervision			-0.039	-0.039		
			[0.053]	[0.053]		
MED_GOV	-3.319	-0.553	-50.355**	-39.332*	-13.884*	-7.808
	[6.742]	[6.021]	[20.092]	[20.486]	[8.172]	[9.465]
POP014	7.462	6.884	-17.950	-20.059	-3.434	-4.418
	[4.890]	[4.551]	[17.139]	[17.381]	[4.681]	[5.041]
POP65P	-0.725	1.083	-7.104	0.304	0.212	4.355
	[7.771]	[8.441]	[21.371]	[22.588]	[10.830]	[10.902]
URB	-1.555	-2.115	-19.082**	-21.495**	2.257	0.777
	[2.597]	[2.751]	[9.525]	[9.034]	[3.775]	[3.463]
ASY	0.487**	0.460**	-0.385	-0.481	-0.097	-0.141
	[0.200]	[0.196]	[0.821]	[0.868]	[0.420]	[0.423]
Y2003	-0.083	-0.096	0.417	0.364	0.09	0.058
	[0.085]	[0.083]	[0.319]	[0.303]	[0.185]	[0.186]
Obs.	248	248	248	248	248	248
Provinces	31	31	31	31	31	31

Notes: Bootstrap standard errors in brackets (the bootstrap is based on 1000 redraws and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

Table 5.9. Fixed effects estimates of economic fluctuations on TB, 2002-2009.

	TB mortality		TB incidence	
	(1)	(2)	(3)	(4)
Δln(GRP per capita)	0.988	1.433	257.309**	88.857
	[0.623]	[1.020]	[108.712]	[120.561]
Δln(GRP per capita)*Inland		-1.129		249.189
		[1.203]		[188.228]
ln(GRP per capita)	0.139	0.088	34.391	18.988
	[0.221]	[0.258]	[29.719]	[26.904]
ln(GRP per capita)*Inland		0.144		25.611**
		[0.138]		[12.929]
MED_GOV	0.291	-0.534	-692.509***	-833.319***
	[2.296]	[1.788]	[226.328]	[229.282]
POP014	-0.226	0.008	-125.726	-114.206
	[1.634]	[1.654]	[244.127]	[255.921]
POP65P	2.783	2.261	160.493	60.017
	[2.727]	[2.764]	[332.378]	[347.845]
URB	0.696	0.811	218.530	262.400**
	[1.034]	[1.022]	[141.648]	[127.655]
ASY	-0.153**	-0.142**	-14.849	-14.364
	[0.070]	[0.066]	[9.439]	[9.211]
Y2003	-0.009	-0.006	-7.016***	-6.083**
	[0.020]	[0.020]	[2.599]	[2.501]
Observations	248	248	248	248
Provinces	31	31	31	31

Notes: Bootstrap standard errors in brackets (the bootstrap is based on 1000 redraws and the replications is based on provinces clusters). *** p<0.01, ** p<0.05, * p<0.1.

5.4. DISCUSSION AND CONCLUSIONS

Reducing maternal and child mortality, improving child malnutrition status and halving TB prevalence are among the key tasks that the MDGs aim to tackle by 2015. Globally, most countries are on a path to achieving these goals; however some less developed nations still lag behind the target. Economic development is a factor that may help sustain the improvement in health status. A number of recent economic crises (e.g. Asian financial crisis in 1997, global financial crisis from 2007 to 2010, Greek debt crisis from 2010) have brought to our attention the potential impact of short-run economic fluctuations on health outcomes. This chapter contributes to the current literature by presenting empirical evidence from mainland China based on both national time series correlation analysis and provincial-level regression analyses. In summary, the results reported in this chapter suggest that economic development benefits the maternal and child health outcomes in China, but not TB incidence rates in inland China. Economic fluctuations are also associated with some of the health outcomes investigated. Specifically, controlling for the level of income, there is a negative association between economic fluctuations and PNM, and a positive association between economic fluctuations and TB incidence rate.

Three limitations of this study should be kept in mind when examining the results. Firstly, although both health outcomes and economic growth have been the subject of several fluctuations for mainland China, economic growth rates are positive for the whole time period considered. In this context, positive shocks to the economy are mainly investigated. The second limitation is related to the issue of causality as discussed in the methodology section. This is a common difficulty facing empirical analyses which use aggregate data. The instrumental variable approach has been adopted for income variables when reverse causality may exist. However, lacking a good instrument for government health expenditure, the causal conclusion between government health expenditure and health outcomes cannot be attained. In future analyses, a good instrument for this should be ideally sought. Thirdly, district-level or county-level data could be used for this analysis; however, only the aggregated provincial/national level data are publicly available for China.

Bearing these limitations in mind, the results of this chapter have mainly supported the previous empirical findings showing wealth is significantly associated with health, while short-term economic fluctuations may have different effects on health outcomes of interest. This finding calls for special policy attention during economic fluctuations for China, as well as potentially being generalisable to other similar developing countries. As identified in several health outcomes equations, the associations between income and health are much stronger in the less developed regions; it is suggested that the improvement of living standards in less developed regions can have a larger effect on health gains. Special programs that target the poor could be established in periods of economic fluctuation.

CHAPTER 6

SUMMARY AND POLICY IMPLICATIONS

6.1. SUMMARY OF FINDINGS AND LIMITATIONS

6.1.1. Summary of Findings

The relationship between health investment, health outcomes and economic growth has been a major focus for both researchers and policy makers. The economic development of China in recent decades has provided an ideal platform for analysing relationships between these factors. Based on publicly available national time series and provincial-level panel data sets, a systematic analysis has been undertaken here to shed new light on these relationships.

Chapter 2 found that for a relatively long-run time period, health investment has exhibited a significantly positive effect on economic output in China. It is important to provide policy-makers with evidence concerning whether or not health investment should be one key issue that needs to be considered when making local development plans. China is a country that has developed unequally across provinces. Considering the regional heterogeneity of mainland China, the results based on provincial-level data in Chapter 3 suggests that the magnitude of health investment on economic output is stronger in inland China (a less developed region), compared to coastal China (a relatively developed region). This result highlights the non-linear effect of health investment on economic output, recently discussed in both theoretical and empirical analyses using cross-country data sets. There is some evidence suggesting the nexus between health investment and economic growth is running in a bi-directional way, where economic output also impacts on health investment. The literature which uses survey data suggests that putting resources into health in relatively poor areas of China (inland or rural China) significantly benefits development.

The health spending behaviour analysis in the short-run, conducted in Chapter 4, provides evidence that within the decentralisation process the provincial government health expenditure and private health expenditure have both exhibited pro-cyclical behaviour alongside economic fluctuations: both public and private health expenditure significantly increase during economic booms. The tax revenue redistribution effect as proxied by central government transformation is significantly related to government health expenditure in inland China. In the time period analysed, although there were economic fluctuations, the Chinese economy has mainly gone through a period of expansion. Thus the results found in this analysis may not reflect the position if the economy undergoes a severe crisis. However, with the limited financial protection effect of medical insurance, this study raises concerns that the population's health status may face a negative impact if the pro-cyclical behaviour remains symmetrical in both good times and bad times.

In the final empirical analysis chapter, this thesis has shown that economic development has significantly impacted on several health indicators in the Millennium Development Goals. Depending on the health outcome indicator, opposing effects have been found: economic growth has reduced the maternal mortality ratio, perinatal mortality rate (PNM) and undernourishment rates among children under 5 years age, while it increased the Tuberculosis (TB) incidence rate in inland China. Controlling for the level of income, this chapter further found that economic fluctuations are significantly negatively associated with PNM while positively associated with TB incidence. The increasing trend of economic output in China will keep benefiting maternal and child health outcomes. However, economic prosperity may also have negative impacts on some public health issues.

6.1.2. Limitations

The limitations of the research in this thesis have been discussed in the relevant chapters. A summary is presented here with a suggested future research agenda.

6.1.2.1. Variable and data issues

- **Health capital proxy –** health capital proxies used in Chapter 2 and 3 are health input (or health resource) variables. The more accurate measurement of the classic health outcome variable adopted in most applied economic analysis, life expectancy at birth, is only available for the years 1990 and 2000 based on census data. Some maternal and child health outcome variables are only publicly available at the provincial level from 2002 onwards. The health input variables are thus utilised following recent literature in this field. It is shown in the literature that health inputs are significant determinants for health outcomes in China. However, how accurately health input variables reflect the health capital stock is unclear. The distribution of health capital between urban and rural China is not considered in this study. As shown in Chapter 5, for several maternal and child health outcomes there exists significant rural-urban disparity. This could be explored further when more detailed data are available in the future.

- **Education capital proxy –** education capital proxies used in Chapter 2 and 3 are raw measurements and not ideal. The chosen proxy is mainly due to data availability. It is common in both empirical macro studies and micro studies using survey data, that the education proxy also suffers from measurement issues. Take the classic proxy, the average schooling years per capita or per worker as an example. In macro analysis, based on limited information, an additional year of education in university is assumed to have the same impact as another one year of education in high school or primary school. This is clearly a strong assumption as recent literature suggests different levels of education have a different impact on economic growth in different development stages. In a micro study, more detailed information about education history can be obtained for each individual. So whether a person has received a Bachelor degree or has a high school diploma can be included as dummy variables directly to control the non-linear effect of education on labour market outcomes. However, there is also an assumption that the quality of education received across regions is the same. This is a major limitation faced by current empirical studies.

- **Health expenditure data –** As discussed in Chapter 4, total health expenditure can be divided into three components; government, social and private health expenditure. In the panel data analysis, both provincial-level government and private health expenditure has been studied. For the part that is left, social health expenditure, only one key component (basic medical insurance expenditure of urban residents) is available for this study.

- **Population data –** There are usually two sources of population statistics provided by the local statistical bureau. The first is based on the number of individuals who have local residence permits (i.e. *hukou*) no matter whether he/she is currently living in that city (household registered population), while the second is based on the number of individuals who actually lived in the city (permanent population). The difference between these is not trivial for some provinces, since the people in inland China tend to temporarily migrate to coastal China for higher incomes. For example, in 2009 the household registered population and the permanent population in Hunan province (inland China) are 70 million vs. 64 million, while in Guangdong province the figures are 101 million vs. 84 million. In this thesis, the population of each province is constructed based on the number of permanent people in each province.

- **Unobserved factors –** Other than variables controlled for in the empirical analysis in this thesis, it is possible that other variables should also considered in explaining the health-growth relationship (e.g. institutions, fiscal decentralisation, etc). These should be considered when data are available.

6.1.2.2. Identification issues

- **Economic fluctuations –** The analyses of economic fluctuations on health spending behaviour in Chapter 4, and health outcomes in Chapter 5, are based on routine fluctuations in Gross Domestic Product (GDP) within different provinces. Although this is common in the applied literature, the results and conclusions attained may be biased when an economy undergoes "severe" crises.

- **Financial protection effect of basic medical insurance –** In Chapter 4 the cyclical movement of private health expenditure of urban residents is regressed on the cyclical movement of basic medical insurance expenditure of urban residents using aggregate level data. It is concluded that the financial protection effect is limited since private health expenditure moves pro-cyclically with basic medical insurance. However in this case, since the population cannot be distinguished between insured and uninsured groups, it is not possible to identify the treatment effect of basic medical insurance *per se*. It would be ideal to use survey data to seek direct answers to this research question.

6.1.2.3. Methodology issues

- **Homogeneous and heterogeneous estimator –** When working with panel data sets, the choice of homogeneous vs. heterogeneous estimator needs to be considered. Currently there is no clear answer to this question. In Chapter 3, while a relatively longer time period is analysed, results from both homogeneous and heterogeneous estimators are reported to seek a robust conclusion. When the time period is relatively short as in Chapter 4 and 5, only a homogeneous estimator is adopted to allow for a greater degree of freedom and increase the efficiency of estimates.

- **Statistical inference –** When dealing with panel data, several assumptions in the error terms need to be tested and handled appropriately to get valid statistical inference. Three assumptions that need special attention in this thesis are: heteroskedasticity, serial correlation and cross-sectional correlation. Take a fixed-effects estimator as an example: calculating the cluster robust standard errors can help handle the heteroskedasticity and serial correlation issues but not the cross-sectional correlation problem. Several methods have been proposed to calculate standard errors that are robust to heteroskedasticity, serial correlation and cross-sectional correlation. The Driscoll-Kraay standard error used in Chapter 3 is an example. However, because it is based on asymptotic theory, to get a better estimation it is not likely to be reliable unless the time period is longer. In Chapter 4 and 5, since the time period is relatively short, a bootstrap standard error that is robust to heteroskedasticity and serial correlation is used.

- **Casual effects and correlations –** In applied macro analysis, it is common to have a problem where the estimates can be casual effects or just correlations. Reverse causality and omitted variable issues commonly confront researchers. The instrument approach is a natural starting point to seek a causal explanation. However it is not always easy to find a perfect instrument in an applied macro study. For example, the instrument for economic growth used in Chapter 4 is the lagged economic growth and/or weather shock. It is commonly adopted in the literature, but it is not perfect. In Chapter 5, fixed asset investment has been used as an instrument for economic output; however, a valid instrument has not been found for government health expenditure. Seeking better instrument is a key point for the research agenda.

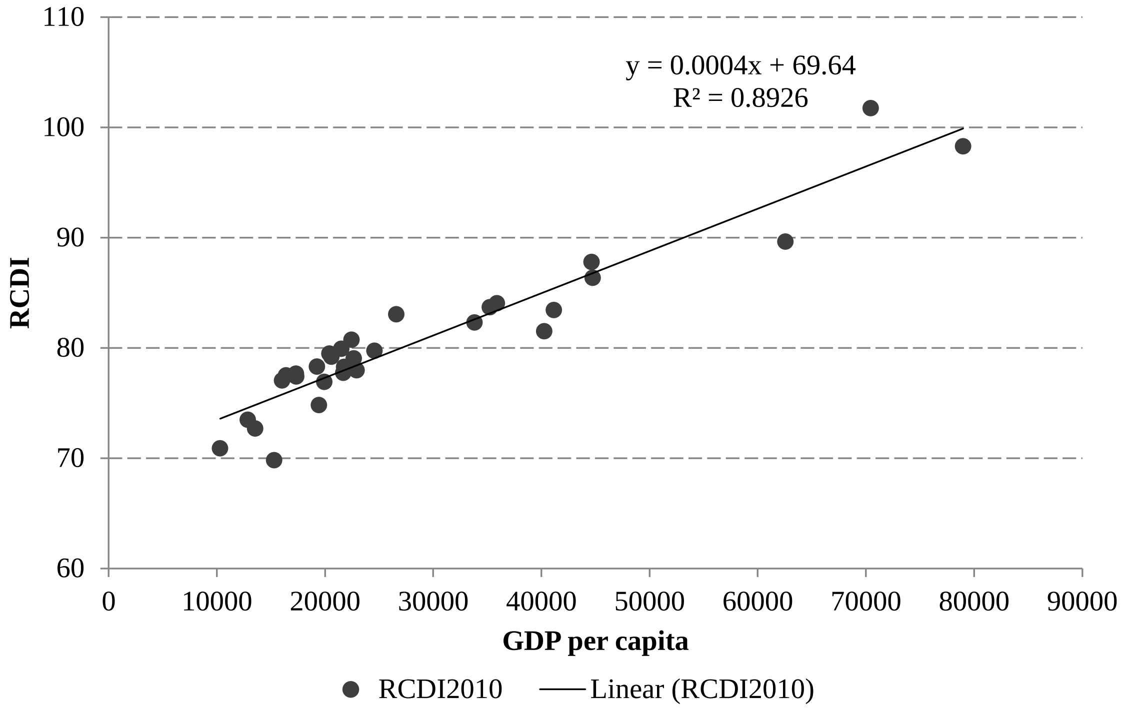
6.2. POLICY IMPLICATIONS AND CONCLUSIONS

6.2.1. Health should be included within an inclusive development strategy

Health issues have attracted less attention from local governments in China since the economic reforms in the early 1980s. It is perhaps not surprising that when political rewards for the governors are tied closely to the GDP growth rate, the incentives for local governors are to start new investment projects rather than provide public services for a given rise in government revenues (Zhou, 2007; Ma and Yi, 2011); health issues are not their top priority, even if health investment can have a positive effect on the economic performance in the long run. The Chinese government has begun to acknowledge the strategic role of health investment in recent years. As reflected in the Chinese government's Eleventh Five-Year Plan for National Economic and Social Development, health has become a priority focus of the government's strategy plan (Tandon, 2006). To shift from a GDP growth focused development to a harmonious development, it would be helpful to have a measurable monitoring system that is valued by the local government.

Several research institutes have worked on a new ranking system to capture multiple dimensions of human development in China. For example, based on the principle of the Human Development Index, the National Survey Research Center at Renmin University of China has developed an index named RUC China Development Index (RCDI)[50] to reflect the multiple dimensions of social, economic and environmental development. The RCDI is composed of four dimensions (or fifteen indicators): health, education, living standards and social environment. Within the health dimension, three indicators are used: life expectancy at birth, infant mortality rate and number of beds in medical facilities per 10,000 population. A simple regression analysis between RCDI and GDP per capita, using data in 2009, is shown in Figure 6.1. As can be seen, there is a positive correlation between RCDI and GDP per capita. There might be some concerns as to whether RCDI is sensitive in measuring the overall development, but the key ideology behind the index, that development is a multi-dimensional concept, should be acknowledged.

[50] See http://www.nsrcruc.org/index/rcdi for details.



Source: RUC China Development Index 2010.

Figure 6.1. Refression between GDP per capita and RCDI, 2009

6.2.2. Central government transition has a redistribution effect on poor areas of China and should be maintained

The cyclical analysis conducted in Chapter 4 suggests that central government transfers are consistently significant and positively related to local government's health expenditure in inland China, but not in coastal China. This finding provides evidence that after fiscal reform in 1994, the redistribution of tax revenue from the central government has been an important channel for funding healthcare in less developed parts of China. Given this, inland China may have an opportunity to catch up with coastal China regarding health resources accumulation and improved health status. However, this also raises some concerns that in the case of an economic downturn, when both central government and local government would be experiencing shrinking tax revenues, health care should still be a key area in investment terms for government, especially in inland China where governments are relatively lacking in financial capacity.

6.2.3. The effectiveness of expanding basic medical insurance coverage should be evaluated

The Chinese government has been successful in expanding basic medical insurance coverage for both rural and urban residents based on a voluntary purchasing principle. However, as has been found in studies using survey data for New Rural Cooperative Medical Scheme in rural China and Urban Resident Basic Medical Insurance in urban China, basic medical insurance schemes have successfully increased the accessibility of health services visits, but the financial protection effect is still limited. This is also supported in this thesis when provincial-level aggregate data are analysed. Besides, there is also evidence suggesting that possibly owing to the low reimbursement rate, basic medical insurance did not affect health status after controlling for the endogeneity issue of insurance enrollment (Yi *et al.*, 2009; Lei and Lin, 2009; Chen and Jin, 2012).

Effective medical insurance coverage has been found to have a protective effect on insured people’s health spending behaviour and health status during an economic crisis in Thailand (Water *et al.*, 2003). For policy-makers, a more important task may be to let insured people actually feel financial relief after purchasing the insurance. Currently local governments are fairly cautious on spending medical insurance funds as shown in Chapter 4. The medical insurance funds are based at city level in urban areas or county level in rural areas. Several drawbacks exist. Firstly, the risk resistance capabilities are limited as the insurance pool is small. Secondly, this may incur higher management costs. Thirdly, it is inconvenient for the insured to transfer between two different insurance schemes when they migrate or get reimbursed in time for the migrant workers. Indeed the benefit packages of basic medical insurance in each city/county are usually set based on experience and were not calculated based on the needs/demand of the local population. All these factors may hamper the effectiveness of basic medical insurance in China.

6.2.4. Economic expansion can have opposite effects on different health indicators

Health status is a multi-dimensional concept. Although the two-way relationship between health and the economy usually suggests that economic expansion will have a positive effect on health status, it also depends on the health outcome measures. The evidence in Chapter 5 shows that higher economic output levels have significantly contributed to the improvement of several maternal and child health outcomes, while on the other hand they also significantly increased the TB incidence rates in inland China. Furthermore, there is some evidence suggesting that after controlling the level of income, economic fluctuations also impact on PNM and TB incidence rates in the short term.

As shown in Chapter 1, broadly speaking the main burden of diseases in China is shifting away from infectious diseases to chronic diseases. However the threat of infectious diseases (e.g. TB), are still a public health concern. Economic prosperity will not eliminate those threats automatically. Economic prosperity and urbanisation may increase the population density and further the likelihood of some infectious disease transition.

6.2.5. Conclusion

The relationship between health investment, health outcomes and economic growth is complex. A considerable body of research in this area uses data from developed countries, while only limited evidence comes from developing countries. This thesis makes an important contribution to the understanding of the above relationship using data from China, one of the largest developing countries in the world. The significant positive relationship of health investment on economic output in the long-run means the health of the population is important for long-term sustainable development of the economy. On the other hand, the short-run cyclical nature of public and private health spending behaviour calls for policy makers' attention, as pro-cyclical spending behaviour may be detrimental for a population's health status in economic downturns. Experience from other Asian countries during the Asian Financial Crisis suggests that medical insurance is one tool that will effectively help maintain quality healthcare during economic downturns, while the current evidence from urban China casts some doubts on its financial protection effect. The multi-dimensional nature of population health status is evidenced by different and sometimes opposite relationships with economic development. Health investment and the acknowledgement of its role on human development are crucial for the developing world.

APPENDIX 1

Appendix 1.1. Comparison of predicted life expectancy at birth

The Figure A.1 plots the life expectancy at birth for China based on data from Banister (1992) and the World Bank. It can be clearly seen that there exists significant difference in the late 1950s and early 1960s. The significant drop in life expectancy shown in the data of Banister (1992) may correspond to the increased crude death rate (which was caused by the Great Leap Forward and the following famine) shown in Figure 1.3.

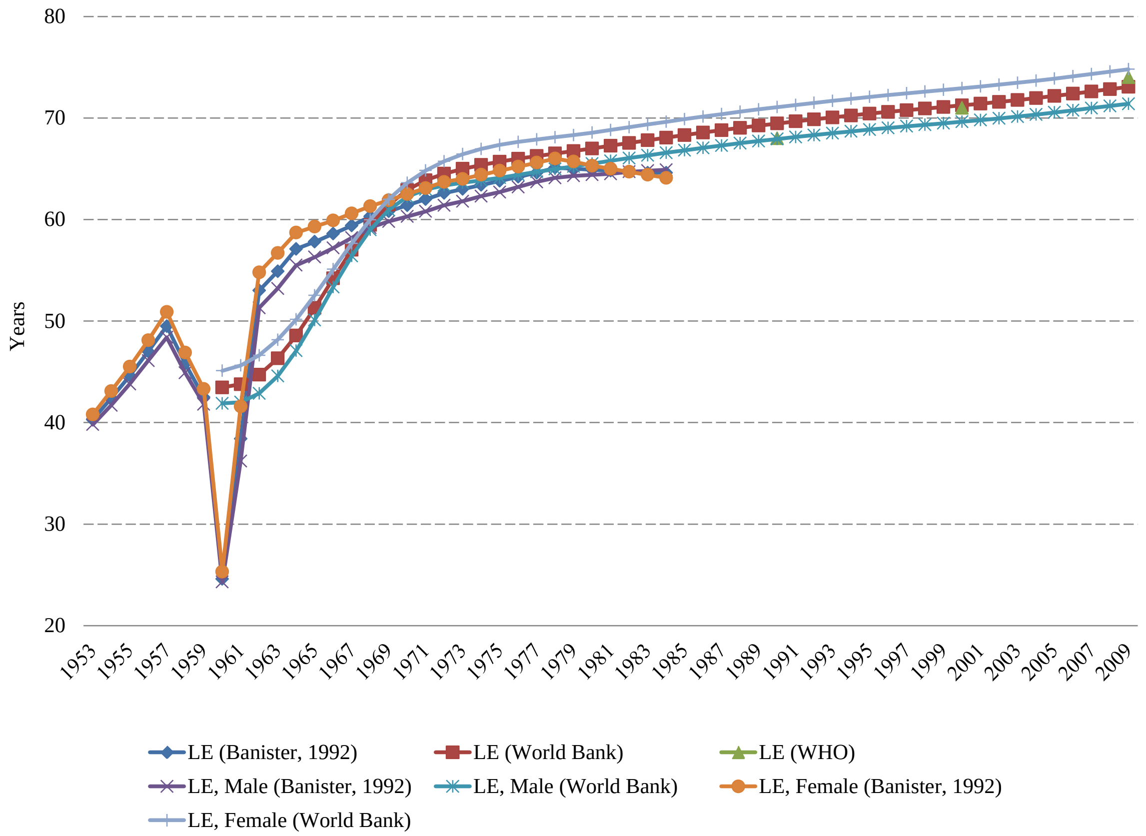


Figure A.1.1. Comparison of life expectancy at birth.

APPENDIX 2

Table A2.1. Literature reviews on the macro-empirical analyses investigating the effect of health on economic growth.

Sources	Dependent variable	Health or health accumulation rate measures / health effect through what channel	Coefficient of health measures (unless specify, all coefficients are significant at the 10% level)	Data / Time period	Econometric methods (main)	Other covariates (please refer to the cited paper for the detail form of covariates)
Knowles and Owen (1995)	change in the natural log Gross Domestic Product (GDP) per working-age person 1960-1985	-ln(80 - life expectancy at birth) / human capital	0.310 ~ 0.381; insignificant in high income countries (Table 1, page 103)	84 countries / 1960-1985	ordinary least squares (OLS) / Two-stage least squares (2SLS)	initial GDP per working-age person, investment/GDP ratio, population growth rate + 0.05, average percentage of the working-age population in secondary school, average years of schooling attained by the population aged 25 and over
Knowles and Owen (1997)	natural log of GDP per working-age person in 1985	-ln(80 - life expectancy at birth) / labour-augmenting factor	0.582 ~ 0.603 (Table 1, page 321)	77 countries / 1960-1985	non-linear least squares (NLS)	initial GDP per working-age person, investment/GDP ratio, population growth rate + 0.05, average percentage of the working-age population in secondary school, average years of schooling attained by the population aged 25 and over

Rivera and Currais (1999a)	change in the natural log GDP per worker 1960-1990	ln(total health expenditure as a percentage of GDP) / human capital	0.21 ~ 0.22 (Table 1, page 264)	24 OECD countries / 1960-1990	OLS	initial GDP per worker, investment/GDP ratio, population growth rate + 0.05 (physical capital depreciation + technical progress), average years of schooling of the labour force
Rivera and Currais (1999b)	change in the natural log GDP per worker 1960-1990	ln(total health expenditure as a percentage of GDP) / human capital	0.22 ~ 0.33 (Table 1, page 763)	OECD countries / 1960-1990	OLS / 2SLS	initial GDP per worker, investment/GDP ratio, population growth rate + 0.05 (physical capital depreciation + technical progress), average years of schooling of the labour force
Arora (2001)	change in the natural log GDP per capita	ln(life expectancy at birth, at ages five, ten, fifteen, or twenty, and stature at adulthood)	0.0269 ~ 1.0276 (Table 2, page 716 & Table 4, page 722)	10 OECD countries / ranges from 1870-1994 to 1891-1994	cointegration	investment/GDP ratio
Bhargava *et al*. (2001)	GDP per capita growth rates 1965-1990 at 5-year intervals	log of lagged adult survival rates (ASR, probability of surviving the 60th birthday after reaching the age 15 years) & interaction between log of lagged ASR and GDP (non-linearity)	0.385 & -0.048 (Table 1, page 431) (significant positive effects of ASR on growth rates until the GDP was approximately 907 in 1985 international dollars)	92 countries / 1965-1990	three-stage least squares (3SLS)	lagged GDP, lagged investment rate, average years of schooling attained by the population aged 15-60, area in the tropics, openness, lagged total fertility rate,

Knowles *et al.* (2002)	average natural log of GDP per worker	-ln(85 - life expectancy at birth) / human capital	0.309; insignificant in less developed countries (Table 1, page 131)	73 countries / 1960-1990	OLS / instrumental variable (IV)-2SLS	investment/GDP ratio, labour force growth rate + 0.05, average years of schooling attained by the population aged 15 and over (male/female), technical efficiency
McDonald and Roberts (2002)	natural log of GDP per capita at the end of each 5-year period	-ln(80 - life expectancy at birth) / human capital	0.106 ~ 0.331; insignificant in OECD countries (Table 1, page 274)	77 countries / 1960-1989	fixed-effects (FE) / pooled OLS	investment/GDP ratio, labour force growth rate + 0.05, mean years of total education, lagged GDP per capita
Webber (2002)	change in the natural log GDP per worker	ln(intake of calories per head)	insignificant	46 countries / 1960-1990	OLS / 2SLS	initial GDP per worker, investment/GDP ratio, labour force growth rate + 0.05, average enrolment in primary education
Bloom *et al.* (2003)	gross domestic saving rate (measured as total savings divided by gross domestic income)	life expectancy at birth / health in general	0.460 (Table 4, page 335)	68 countries / 1960-1997	2SLS	growth rate of GDP per capita during preceding decade, age structure
Rivera and Currais (2003)	change in the natural log GDP per worker 1960-2000	ln(total health expenditure as a percentage of GDP) / human capital	0.20 ~ 0.26 (Table 3, page 320)	OECD countries / 1960-2000	OLS / 2SLS	initial GDP per worker, investment/GDP ratio, population growth rate + 0.05, average years of schooling attained by the population aged 25 and over
Bloom *et al.* (2004)	change in the log GDP per capita 1970-80 and 1980-90	life expectancy at birth / human capital	0.040 (Table 4, page 10)	104 countries / 1960-1990	non-linear two-stage least squares (NL2SLS)	physical capital, labour force, average years of schooling, average work experience of the work force and its squared term

Gyimah-Brempong and Wilson (2004)	annual growth rate of GDP per capita averaging over 4-year intervals	(1) total health expenditure as a percentage of GDP or government health expenditure as a percentage of GDP (investment in health capital), (2) inverse of the child mortality rate per 1000 births & squared term (health capital stock)	(1) 0.0907 (Africa) & 0.0493 (OECD), (2) 0.3384 & -0.1420 (Africa) & 0.0591 & -0.0012 (OECD)	21 Sub-Saharan African countries or 22 OECD countries / 1975-1994 or 1961-1995	Arellano-Bond generalised method of moments (GMM)	lagged GDP per capita, investment/GDP ratio, average years of schooling attained by the population aged 15 and over, openness, political instability, % population that is 14 years or younger, % of population that is 65 years or older
Rivera and Currais (2004)	change in the natural log gross value added (GVA) per worker	(1) ln(total government health expenditure), (2) ln(public health consumption expenditure), (3) ln(public health investment expenditure) / human capital	(1) 0.13 ~ 0.15; (2) 0.15 ~ 0.16; (3) insignificant for public health investment expenditure (Table 2, page 878)	17 Spanish regions / 1973-1993	OLS	initial GVA per worker, investment/GVA ratio, working age population growth rate + 0.05, % of pop with secondary school qualifications

Alsan *et al*. (2006)	log gross Foreign Direct Investment (FDI) inflows	life expectancy at birth / human capital	0.065 ~ 0.075 for full sample (Table 3, page 622), 0.089 ~ 0.093 for low- and middle-income country sample (Table 4, page 624), insignificant for high-income country sample	74 countries / 1980-1990 & 1990-2000	N/A	initial population, initial GDP per capita, initial openness, % of pop >= 25 have completed secondary schooling, bureaucratic quality, corruption governance, telephones mainlines per 1,000 capita, distance to major markets, landlocked, pop coastal (%)
Cole and Neumayer (2006)	Total factor productivity (TFP)	(1) ln(undernourishment), (2) ln(malaria incidence) (3) ln(access to safe water) / TFP	(1) -0.17 ~ -0.21 (Table 3, page 928); (2) -0.34 ~ -1.06 (Table 4, page 930); (3) -0.09 ~ -0.63 (Table 5, page 931)	52 countries / (1) 1980, 1991 & 1996; (2) 1996 & 1994; (3) 1970, 1975, 1985, 1988 & 1993	FE / IV- 2SLS	share of trade in GNP, share of agricultural value added in total GNP, inflation rate
McDonald and Roberts (2006)	natural log of GDP per capita at the end of each 5-year period	predicted ln(infant mortality rate) / human capital	0.631 ~ 0.966 (Table 2, page 240)	112 countries / 1960-1998	Arellano-Bond GMM / IV-2SLS	investment/GDP ratio, population growth rate + 0.05, secondary school enrolment rate, lagged GDP per capita

Beraldo *et al*. (2009)	change per year in the natural log of GDP per capita	(1) change per year in the ln(total health expenditure per capita), (2) change per year in the ln(public health expenditure per capita), (3) change per year in the ln(private health expenditure per capita) / labour productivity	(1) 0.1042, (2) 0.0697 ~ 0.0720, (3) 0.0081 ~ 0.0095 (Table 6, page 953, static IV estimates)	19 OECD countries / 1971-1998	Arellano-Bond GMM / 2-way random effects / 2-way FE	investment/GDP ratio, private capital stock, public and private education expenditure per capita, Gini index, three political proxies, fiscal decentralisation
Li and Huang (2009)	natural log of GDP per capita	(1) ln(number of doctors per 10,000 persons), (2) ln(number of hospital beds per 10,000 persons) / human capital	(1) 0.18 ~ 0.20 (Table 15, page 384, 2SLS estimates), (2) 0.26 ~ 0.55 (Table 15, page 384, 2SLS estimates)	28 provinces of China / 1978-2005	FE / 2SLS	lagged GDP per capita, investment/GDP ratio, % of people with secondary or above education, student to teacher ratio
Astorga (2010)	growth of GDP per capita	lagged life expectancy at birth / human capital	1.13 (Table 1, page 236)	Latin America / 1900-2004	3SLS / seemingly unrelated regression (SUR)	investment, openness, terms of trade, lagged illiteracy rate,

Benos and Karagiannis (2010)	natural log of GDP per capita	ln(number of medical doctors per 1000 inhabitants) / human capital	0.092 (Table 7A.5, page 156), 0.055 ~ 0.058 for high-income regions (Table 7A.6, page 157) vs. 0.087 ~ 0.089 for low-income regions (Table 7A.7, page 158)	51 regions of Greece / 1981-2003	Arellano-Bond GMM	initial GDP per capita, public investment, population growth rate, number of students attending the lower / upper secondary level of education, student-teacher ratio
Bloom *et al*. (2010)	growth rate of GDP per capita	life expectancy at birth / human capital	0.092 ~ 0.106 (Table 5, page 26, 2SLS estimates)	cross-countries / 1960-2000	2SLS	initial GDP per capita, investment/GDP ratio, openness, average years of schooling attained by the population aged 15 and over, bureaucratic quality, tropical area, sectoral change, share of working-age population, growth of share of working-age population
Hartwig (2010)	growth of GDP per capita (5-year averages)	(1) lagged growth of health expenditure per capita (5-year averages), (2) lagged change in life expectancy at birth over 5-year intervals / human capital	insignificant based on 2-step GMM estimator (Column 3, Table 4, page 319 & Column 3, Table 9, page 322)	21 OECD countries / 1970-2005	Arellano-Bond GMM	none

Jamison *et al*. (2010)	natural log of average GDP per capita over a 5-year period	ln(male survival rate) / human capital	0.50 (Column 10, Table 4, page 30)	53 countries / 1965-1990	hierarchical linear modelling (HLM)	physical capital per capita, average years of schooling attained by the population aged 15 and over, total fertility rate, tropics; technical progress determinants equation (coastal, openness)
Narayan *et al*. (2010)	log of GDP per capita	log(health expenditure as a percentage of GDP) / human capital	0.1679 ~ 0.2677 (Table 3, page 409)	5 Asian countries / 1974-2007	panel cointegration (DOLS)	gross fixed capital formation/GDP ratio, exports/GDP ratio, imports/GDP ratio, R&D/GDP ratio, education expenditure/GDP
Suhrcke and Urban (2010)	GDP per capita	cardiovascular disease mortality rates per 1000 inhabitants in working age (CVD) & interaction between CVD and lagged GDP per capita (non-linearity) / human capital	0.441 & -0.047 (Column 4, Table 3, page 1487) (a positive (negative) impact of CVD on growth below (above) a threshold level of income below 12,255 PPP US $); when splitting sample according to the threshold, the negative effect of CVD is only significant in the high-income countries	61 countries / 1960-2000	Blundell-Bond GMM	lagged GDP per capita, investment/GDP ratio, openness, average years of schooling, fertility rate, adult mortality rate

Anderson *et al*. (2011)	log of GDP per capita or log of GDP per worker	(1) log(Ultraviolet (UV) radiation, an index of Erythemal exposure), (2) log(cataract prevalence)	(1) -0.72 ~ -1.39 (Table 1 & 2, pages 36-37), (2) -0.28 ~ -0.32 (Table 1 & 2, pages 36-37)	1990-2000	OLS	latitude, mean elevation, temperature, precipitation, country area, distance to coastal, distance to river, year of Neolithic Transition, a set of continent dummies
Aghion (2011)	change per year in the log GDP per capita 1960-2000, multiplied by 100	(1) change per year in the log life expectancy at birth, multiplied by 100, (2) initial log(life expectancy at birth)	(1) 172.58 ~ 184.48, (2) 8.10 ~ 9.10 (Table 3, page 10)	96 countries / 1960-2000	2SLS	initial GDP per capita
Swift (2011)	(1) total GDP, (2) GDP per capita	life expectancy at birth / human capital	(1) 2.679 ~ 9.286 (Table 4, page 314), (2) 2.397 ~ 7.432 (Table 5, page 315)	13 OECD countries / ranges from 1831-2001 to 1921-2001	cointegration	none
Audibert *et al*. (2012)	annual growth of GDP per capita 2000-2004	variants of Disability-adjusted life year (DALYs) per capita / human capital	-0.108 ~ -0.365 (Table 1, page 15)	153 countries / 2000-2004	IV-GMM	initial GDP per capita, investment/GDP ratio, population growth rate, government consumption ratio to GDP, openness, inflation rate, school enrolment lagged, institutions

Notes: A detail review of estimates of the effect of health on economic growth by Barro *et al*., Bloom *et al*. and several other authors' papers published from 1994 to 2000 could be referred to Table 1 at Bloom *et al*. (2004, pages 2-4). Audibert *et al*. (2012, pages 29-32) also provides a detailed summary of empirical studies published between 1994 and 2007 in this theme.

APPENDIX 3

Appendix 3.1. The prediction of average schooling years per worker

In the first step, the average years of schooling (AYS_{it}) per worker is calculated for each region based on available data from the *China Labor Statistical Yearbooks* since 1996.

Extending Equation (2.6) in Chapter 2 to a panel setting, the following equation has been adopted:

$$AYS_{it} = 1E_{1it} + 5E_{2it} + 8E_{3it} + 11.5E_{4it} + 14.5E_{5it} \tag{A3.1}$$

where E_1 to E_5 stands for the different proportions of illiterate and semiliterate, elementary school, junior high school, senior high school and, college and above in time periods t. The lengths of schooling cycles for the above categories are assumed to be 1, 5, 8, 11.5 and 14.5 respectively.

In the second step, provincial-level data are used to regress average schooling years per worker on a set of control variables where both data are available (1996-2006), and then predict average schooling years per worker for the whole time period (1978-2006). The control variables include: the proportion of labor force in an urban area (z_1), the proportion of labor force in the primary industry (z_2), the proportion of labor force in the secondary industry (z_3), the proportion of the labor force that is male (z_4), the number of students enrolled in college and above per 10,000 employed persons (z_5), the number of students enrolled in secondary school per 10,000 employed persons (z_6), the number of students enrolled in primary school per 10,000 employed persons (z_7), the teacher-student ratio for secondary school (z_8), the teacher-student ratio for primary school (z_9), the dummy variable for coastal sample (z_{10}). Using a random-effects technique, the following model is estimated:

$$AYS_{it} = \zeta_1 z_{1it} + \zeta_2 z_{2it} + \zeta_3 z_{3it} + \zeta_4 z_{4it} + \zeta_5 z_{5it} + \zeta_6 z_{6it} + \zeta_7 z_{7it} + \zeta_8 z_{8it} + \zeta_9 z_{9it} + \zeta_{10} z_{10i} + u_{it} \quad \text{(A3.2)}$$

The estimated result is:

$$\begin{aligned} AYS_{it} = {} & \underset{[0.62]^{***}}{1.93} * \hat{z}_{1it} - \underset{[0.93]^{***}}{2.90} * \hat{z}_{2it} - \underset{[1.01]}{1.30} * \hat{z}_{3it} + \underset{[5.79]}{5.28} * \hat{z}_{4it} + \underset{[0.00043]}{0.00060} * \hat{z}_{5it} + \underset{[0.00019]^{***}}{0.00074} * \hat{z}_{6it} \\ & - \underset{[0.000095]}{0.00015} * \hat{z}_{7it} - \underset{[1.29]^{***}}{3.58} * \hat{z}_{8it} + \underset{[6.45]}{4.08} * \hat{z}_{9it} + \underset{[0.18]^{***}}{0.48} * \hat{z}_{10i} + \underset{[3.16]}{4.85} \end{aligned} \quad \text{(A3.3)}$$

where standard errors are reported below the coefficients, *, **, and *** stand for significance at the 10%, 5%, and 1% level respectively. The calculated within, between, and overall R squared are 0.63, 0.80, and 0.77 respectively. The simple correlation between ln(education capital per worker) used in the main analysis and the predicted ln(average years of schooling per worker) is 0.450.

APPENDIX 4

Appendix 4.1. The multivariate Beveridge-Nelson decomposition technique

In this appendix, a brief introduction of the multivariate Beveridge-Nelson decomposition approach is summarised; the content presented below is mainly derived from Garratt *et al*. (2006a) and Garratt *et al*. (2006b).

Consider a random variable y_t, which can be decomposed into two components, a permanent series and a cyclical series, i.e. $y_t = y_t^P + y_t^C$. The cyclical component is the interest of this study and it is usually calculated as $y_t^C = y_t - y_t^P$.

Several techniques can be used to estimate the permanent component. For example, if adopting one of the most widely used methods, the Hodrick-Prescott filter, the permanent component can be determined from solving the following problem:

$$\min \frac{1}{T}\sum_{t=1}^{T}(y_t - y_t^P)^2 + \frac{\lambda}{T}\sum_{t=2}^{T-1}[(y_{t+1}^P - y_t^P) - (y_t^P - y_{t-1}^P)]^2 \qquad \text{(A4.1)}$$

where λ is a non-negative number and T is the number of usable observations. In this minimization problem, λ is specified as an arbitrary number reflecting the penalty of incorporating cycles into the trend. If annual data is used, λ is usually recommended to be 6.25. For a set of variables, one simply adopts the same technique to each variable to extract the permanent components. As can be seen later, compared with the multivariate decomposition technique, this univariate approach will omit any long-run restrictions or short-run interactions that might exist among a group of cointegrated series.

Now consider an m × 1 vector of random variables $z_t = (y'_t, x'_t)'$ and take any arbitrary partitioning of z_t into permanent and cyclical components, i.e. $z_t = z_t^P + z_t^C$. The permanent component can be further divided into two parts, the deterministic and the stochastic components, i.e. $z_t^P = z_{dt}^P + z_{st}^P$. z_{dt}^P and z_{st}^P are defined as:

$$z_{dt}^P = g_0 + gt \tag{A4.2}$$

$$z_{st}^P = \lim_{h\to\infty} E_t(z_{t+h} - z_{d,t+h}^P) = \lim_{h\to\infty} E_t[z_{t+h} - g_0 - g(t+h)] \tag{A4.3}$$

where $E_t(\cdot)$ denotes the expectation operator conditional on the information at time t, taken to be $\{z_t, z_{t-1}, \ldots, z_0\}$; g_0 is an m × 1 vector of fixed intercepts; g is an m × 1 vector of (restricted) trend growth rates; t is a deterministic trend term. Since the cyclical components must satisfy $\lim_{h\to\infty} E_t(z_{t+h}^C) = 0$, then it must follow that:

$$z_t^P = \lim_{h\to\infty} E_t(z_{t+h} - gh) \tag{A4.4}$$

Equation (A4.4) is the basis of permanent/cyclical decomposition of z_t. Next, one specifies the suitable vector error correction model for z_t and under which the permanent/cyclical properties of x_t would have a direct bearing on those of y_t.

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