UPPSALA UNIVERSITET

Economic Studies 126
Department of Economics

Essays on Development, Institutions and Gender

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Department of Economics, Uppsala University

Visiting address: Kyrkogårdsgatan 10, Uppsala, Sweden
Postal address: Box 513, SE-751 20 Uppsala, Sweden
Telephone: +46 18 471 11 06
Telefax: +46 18 471 14 78
Internet: http://www.nek.uu.se/

ECONOMICS AT UPPSALA UNIVERSITY

The Department of Economics at Uppsala University has a long history. The first chair in Economics in the Nordic countries was instituted at Uppsala University in 1741.

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Abstract

Dissertation at Uppsala University to be publicly examined in Hörsal 2, Ekonomikum, Friday January 21, 2011 at 10.00 a.m. for the degree of Doctor of Philosophy. The examination will be conducted in English. RICKNE, Johanna, 2011, Essays in Development, Gender and Institutions, Department of Economics, Uppsala University, Economic Studies 126, 139 pp. ISBN: 978-91-85519-33-0, ISSN: 0283-7668 urn:nbn:se:uu:diva-137502 (http://urn.kb.se/resolve?urn=urn:nbn: se:uu:diva-137502)

This thesis consists of four self-contained essays.

Essay 1: (with Johan Lyhagen) In the first essay, we develop an improved test of economic convergence or divergence using time series methods. The usefulness of the method is illustrated in an analysis of the growth pattern between Chinese regions in 1952-2007. Comparing all combinations of regional pairs, we find evidence of economic divergence in roughly half of the cases. In the other half, we instead find that regions have grown while maintaining stable income differences.

Essay 2: The second essay compares average earnings and productivities of men and women employed in roughly 250,000 Chinese industrial enterprises. Women are found to earn, on average, 12% less than men. This gender wage gap stems entirely from a wage disadvantage for women with less than twelve years of education. For women with more than twelve years of education, the average wage exceeds the average wage of skilled men. The paper also shows that including firms' financial contributions to social insurance only marginally affects the gender distribution of work compensation. With respect to discrimination, women's earnings disadvantage is entirely accounted for by the gender gap in productivity, indicating that there is no evidence of negative wage discrimination against women.

Essay 3: The third essay takes as a starting point the situation that ten years after the establishment of China's Unemployment Insurance (UI) program, coverage is still incomplete owing to large-scale firm evasion. This study draws on a highly representative panel of Chinese firms for 2001-2005 to

describe the pattern of firm participation across time, ownership, region, and industrial dimensions. A simple theoretical model is then used to derive the prediction that firms are more likely to provide UI if they operate in tighter labor markets. This prediction receives empirical support, but the effect is quantitatively small. Factors other than the labor market situation, such as ownership type, the firm's cost profile, and worker demand, are more important for explaining a firm's provision of insurance.

Essay 4: The fourth essay finds political and legal institutions affect the extent to which the real exchange rates of oil-exporting countries co-move with the oil price. In a simple theoretical model, strong institutions insulate real exchange rates from oil price volatility by generating a smooth pattern of fiscal spending of oil revenue over the price cycle. Empirical tests on a panel of 33 oil-exporting countries provide evidence that countries with high bureaucratic quality and strong and impartial legal systems have real exchange rates that co-move less with the oil price.

To my parents

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Stockholm 24/11, 2010
Johanna Rickne

Contents

Introduction

Around the world, countries continue to struggle for economic development. In the past twenty years, Western-led policies of economic liberalization failed to produce the intended effects in Latin America and Sub-Saharan Africa (Rodrik 2006). In former socialist economies, big bang style economic transitions caused dramatic drops in economic activity. Despite this disappointing record, global poverty decreased more rapidly during this period than ever before. In the most populous country in the world, poverty dropped from 53% to 8% over the 1980s and 1990s (Ravallion and Chen, 2007). After yet another decade of "miraculous" growth, China surpassed Japan as the world's second largest economy.

There are disagreeing views on the nature of China's growth experience. Some consider it a unique policy experiment with the potential of offering valuable lessons to other developing countries. Others are more skeptical and emphasize the country's internal and external instability. Internally, income inequality has skyrocketed, between regions as well as within workplaces. Externally, the act of dismantling the communist welfare system for urban residents led to increased income insecurity which made citizens save more and consume less. Rather than relying on the weak domestic market, the Chinese economy grew by producing export goods for the world market (Lardy, 2007: Prasad, 2008). As a way to reverse the imbalance in trade and improve the welfare of workers, China is currently constructing a highly ambitious social insurance system for its urban population. But progress with expanding insurance coverage has been slow with many workplaces evading their obligations.

Despite the obvious importance of China on the world's economic and political stage, it has been the subject of comparatively little research in the field of Economics. This thesis examines some of the top concerns to developing and transition economies and focuses on the Chinese economy as a case study. Contributions to existing research is made both by improving on statistical tools, using more comprehensive datasets, and exploring new and unanswered research questions.

Increasing income differences is an expected consequence of an economic transition from a planned to a market economic system. In this regard, the Chinese experience could however also be seen as an example of a global trend toward inequality. Regardless, the observation contradicts the key prediction from neoclassical economic theory that incomes should grow more similar over time (Solow, 1956). In essay 1, I and Johan Lyhagen present an

improved time series test for the economic convergence hypothesis. We then demonstrate the relevance and usefulness of our test using growth data for China's regions. With our test, we can move beyond answering the question of whether incomes are already equal. Instead we can investigate how inequality has developed over time, how fast it has evolved, and if it has been reversed in latter years.

The second essay addresses a specific component of China's emerging income inequality, namely that between men and women. By using a uniquely comprehensive firm-level dataset, the essay expands the insights from previous studies in several dimensions. It provides highly representative results about how the gender wage gap varies across firms with state or foreign ownership, industrial sectors, geographical regions, or for workers with different education levels. The existence of gender wage discrimination is assessed by examining whether women receive wages that match their contributions to firm productivity.

The third essay represents the first quantitative study of the determinants of Chinese firms' participation in one of the country's newly reformed social insurance programs. It explores firm provision of unemployment insurance in a five-year data panel. Focus is on whether firms respond to local labor market conditions by providing insurance, but the empirical investigation gives a wider view of the pattern of participation according to institutional factors, the firms' cost profile and workforce characteristics.

The fourth and final essay deals with an economic challenge often named as a main contributor to Russia's failed transition, namely how countries that export natural resources can face problems in developing their industrial sectors because of macroeconomic instability. This essay's main contribution is in the field of empirical macroeconomics. Using data for oil-exporting countries, I test the hypothesis that legal and political institutions can explain cross-country differences in the way that the values of countries' currencies respond to variation in the oil price.

Economic reform and income inequality in China

A main concern, both for Chinese policy makers and citizens, is that the economic development of the past three decades has been highly unequal. Some regions have become a lot richer, while poorer locations have failed to catch up. According to standard measures of income inequality, communist China is now on par with Mexico and Brazil, two of the most unequal countries in the world. During the last decade, the government has aimed at bridging the development gap between the country's Eastern and Western areas by re-directing investment toward low income regions in the West.

Mainstream economic theory states that for given endowments, poor economies should catch up with richer ones (Solow, 1956). In other words, economies with similar growth potential, for example through access to nat-

ural resources or an educated labor force, should eventually arrive at the same level of income. Much research has been dedicated to testing if time makes incomes grow more similar between countries, or between regions within countries (Islam, 2003). The conclusions have varied greatly, with some studies arguing that convergence is happening, and others arguing that it is not. A potential reason for this discrepancy is the use of inappropriate statistical methods.

In **essay 1**, I and Johan Lyhagen propose a new way to test the hypothesis of economic convergence. We argue that the most commonly used time series test of this question is inadequate: it essentially answers the question of whether two regions are already at the same income level, and not the question of whether the income difference between the regions is becoming smaller.[1] In our proposed test, we instead capture how the difference in income between two regions has changed over time, and how fast these changes have taken place. In technical terms, we allow for nonlinear elements in the cointegrating relationships between regional income per capita time series. With this method, we can not only give a more correct answer to the main question of decreasing income differences, we can also examine several new issues such as the timing and speed of the developments in inequality.

We apply our statistical test to China's regions (22 provinces, 5 autonomous regions and 4 self-governing municipalities)[2] over the 1952–2007 period. This dataset represents an appropriate test bed for our proposed method because Chinese regions are larger than most countries in the world, but share a common currency, language and institutional setting. We also know that incomes have become more unequal during the market transition period. A method that can assess the timing and speed of this divergence, and whether it has been reversed in recent years, is therefore a suitable tool for our analysis of this economy.

In the essay, we first establish that our test is highly relevant for examining economic convergence by showing that the nonlinear trend specifications capture more of the growth dynamics between regions than the standard linear form. We then make a number of observations about the development of income inequalities between China's regions over time. The main finding is that incomes have become more unequal, or maintained their level of inequality over the entire period of study. By comparing each region to all the other regions, pair by pair, we conclude that in half of the cases, the difference in average income per capita has remained roughly constant. In the other half of the cases, the income gap has grown. In these cases, inequality has started to grow at the outset of the market reform period in the late

[1] Following proposition 5 in Bernard and Durlauf (1996), the standard time series test of economic convergence requires that the difference between the output series of two regions is stationary and has a zero mean (Bernard and Durlauf, 1996; Evans, 1998)

[2] These administrative areas are directly subordinate to the Chinese central government and correspond roughly to US States.

1970s, and accelerated in the period thereafter. We also reject that this trend of increasing inequality has been reversed in recent years.

How can we interpret the fact that our results are inconsistent with the prediction from neoclassical economic theory? One interpretation could be that that the economic links between China's regions do not meet the basic theoretical preconditions. Most importantly, capital and labor may not move around freely and equalize factor prices. Bank loans are channeled by State-owned banks to specific industrial sectors and localities, and people are retained from moving around freely by the family registration system.[3] If this interpretation were correct, it would imply that increased factor mobility would facilitate a reduction of China's regional inequality.

Although interesting, the big picture of regional inequality does not provide insights into the microeconomic dimensions of the effect of market liberalization on the wage differences between Chinese workers. **Essay 2** addresses one of the most commonly researched dimensions of wage inequality, namely that between women and men. This question is highly relevant in the Chinese case. First, it is worth noting that the female labor force in China is one fifth of the entire female labor force in the world.

We also have reasons to expect that market liberalization may have had different effects for women and men. During the planned economy, China had a strong commitment to women's rights. The gender wage gap was narrow, and a large proportion of all women held full-time jobs (Croll, 1995). With reform, workers were no longer allocated to specific workplaces by the authorities, and firms were allowed to set wages more freely. When employers became allowed to choose between workers and set wages according to their own tastes, they got the power to discriminate in order to satisfy their potential preferences for specific worker types (Becker, 1971). In China, reforms thus could have allowed more room for (male) employers to express a preference for male workers in accordance with the Confucian tradition.[4] Less negative effects could however be expected if we would instead consider the rapid process of internationalization of Chinese workplace in the transition period. As the experience of other emerging economies has shown, women may be the preferred choice of labor in factories producing light manufacturing goods for the international market (e.g. Standing, 1999).

As concluded in Shen and Deng's (2008) review article, studies of gender wage discrimination in China have generally lacked access to representative datasets of high quality. Most studies have used wage information on the individual or household level, which is less appropriate than firm level data

[3] An individual's family registration status (*hukou*) functions like an internal passport system and rarely admits movements of individuals between rural and urban areas, or between cities.

[4] The Confucian ideology emphasizes male-to-male relationships and traditionally assigns a lower status in the hierarchy for women compared to men. The roles of the husband and wife separated, with the wife in a subordinate position and the man as the family's breadwinner.

to examine the question of discrimination.[5] In the few papers that have used observations of firms, the datasets have been small. In the paper most closely related to the thesis chapter, Dong and Zhang (2009) analyze information from less than 1,000 firms. In comparison, the essay included in this thesis uses information covering almost a quarter of a million firms. Together, these firms produced more than 90% of China's total industrial goods in year 2005, meaning that the conclusions have high relevance for China's industrial sector as a whole.

The essay belongs to a group of papers that employ Hellerstein and Neumark's (1999) method of analyzing discriminations using firm-level data. This method is based on comparing not only women and men's wages, but also their productivity. For example, a finding that women earn 20% less than men, but are 30% less productive would be interpreted as there being no evidence of discrimination against female workers.

My results give a number of insights into how China's female industrial workers' wages compare to those of their male colleagues, both depending on education level, if they work in the state or private sector, or which types of products their firms produce. Although not all results can be described in detail in this introduction, a couple of key points deserve mentioning.

For the role of education, my results support those of previous studies in finding that among workers with less than 12 years of education, women's wages are lower than men's, but among workers with more than 12 years of education, women's wages are instead higher. However, in contrast to previous studies, my results also show that the wage advantage of highly educated women is limited to the State-owned sector.

Calculating an average gender wage gap for all women compared to all men (that is, for all levels of education) shows that the average woman's wage is 12% lower than the average man's. Compared to studies of other countries, this is a very narrow gender wage gap. I also conclude that non-wage earnings, in the form of social insurance contributions made on part of the worker by his or her employer, do not contribute to a wider gap in total earnings. An interesting finding regarding the international comparison of gender wage gaps is that the results in the essay do not support that China's narrow gap is due to more equal wage-setting in state-owned workplaces. In fact, the average gender wage gap is found to be narrower in privately owned firms than in state-owned ones.

Substantial variation in female employment and relative wages is also found across industrial sectors. The data shows that the majority of the women are hired in factories that produce light products, such as toys or textiles: in the average textile firm, 66% of the labor force is female, and in the average tailoring firm, that number is 73%. Notably, in these firms, women's wages are more equal to men's than in industries that produce

[5] The most commonly used household level datasets also suffer from under-sampling of private and foreign-owned workplaces.

heavy industrial goods. Although wages do not reflect all dimensions of a job, these results appear to indicate that women have benefitted from the reform period's expansion of light manufacturing sectors, both in terms of employment and relative wages.

In many ways, the results in the essay are different from those previously found for other transition economies when examining the effect of wage liberalization on women's labor market situation. China does not appear to have followed in the tracks of the Central and Eastern European Countries, where researchers have observed a fall in women's labor force participation and a reduction of the gender-wage gap. Nor do the results mirror the maintained female participation rate and lower relative wages of the Former Soviet Union.[6]

Rather, China's experience appears more similar to those of other emerging economies with a rapidly globalizing industrial sector. When comparing women's productivity to their wages, there is no evidence of discrimination. Instead, the results indicate that employers in the recently created industrial sectors have preferred to hire women over men. This observation is supported by qualitative studies of the production system in Chinese factories, where managers often express a preference for female assembly line workers because of their perceived docility and "nimble fingers" (i.e. Pun, 2005).

In sum, the results of the essay do not support the role of gender inequality as a main cause for concern when considering the overall growth of inequality in China. In addition to the gender wage gap for Chinese industrial workers being narrow compared to other countries, women's wages are also found to be more equal to men's in more recently established industrial sectors, and in foreign-owned firms.

More attention should perhaps be paid to the result that women's average productivity is lower than men's. In this regard, the essay shows a wider productivity gap in China compared to what previous studies have found in other countries.[7] This finding highlights the potential of raising overall industrial productivity by finding ways to eliminate structures that contribute to women's productivity disadvantage. Identifying these structures is an interesting topic for future research.

[6] For a literature review that describes this general patterns, see Pastore and Verashchagina (2007).

[7] This finding is standard in the literature, also when using employer-employee data with employee controls and when using advanced controls for labor force changes in firms (e.g. Hellerstein and Neumark, 2005; Barcolucci and Alberto, 2010). In my paper, the existence of a gap is robust to estimation for a pooled firm sample over both 2004 and 2005 using firm fixed effects.

Social (in)security and the Chinese labor market

Under the planned economy, workers in Chinese cities received medical care, pensions and housing from their state-controlled work units. When the state sector was reformed, the unitary social security scheme was replaced by individual insurance programs for health care, pensions etc., in which firms in all ownership sectors were obliged to participate. Because the insurance requirements were not based on laws, but on a patchwork of regulations, many firms evaded their responsibilities, leading to a slow expansion of coverage.[8] In **essay 3** I investigate the mechanisms behind Chinese firms' participation in one of the newly created social insurance programs, namely the Unemployment Insurance (UI) which was set up in 1999.

Before I briefly introduce the results of the essay, it is worthwhile to examine the close connections between the reformed social insurances and the broader picture of China's role in the global economy. For Chinese workers, firm evasion of their unemployment, pension, and health care mandates, implied a dramatic increase in income insecurity. To protect themselves from the risk of joblessness, sickness and old age, workers radically increased their savings and, thus, decreased their consumption.

One of the hot topics of the global economic debate is the issue of the global trade imbalances. This term relates to the fact that some countries are lending money to finance the debt of other countries, while keeping large trade surpluses vis-à-vis these economies. By many, this situation is perceived as unsustainable and as increasing the risk in the global economic system. The high level of private savings in countries like China, directly contributes to the imbalances as these savings are channeled into international lending. The social insurance programs are a key factor in this equation. Along this line of reasoning, the expansion of China's social insurances can be crucial in reducing China's reliance on foreign demand and work as a remedy for the ill of the global imbalances.

Although understanding the determinants of firm provision of social insurance is key to the study of China's future economic development, there is only one previous paper that has examined this issue. In general, both for China and for other countries where social insurance participation is optional, such as the US, gaining access to reliable data on the firm level is a major obstacle. Compared to Nielsen and Smyth's (2007) study of 5,000 Shanghainese firms, selected by the authorities, my essay explores a panel of firms that together account for 70-95% of China's total value added in 2001-2005.

[8] Arguably, the slow expansion of coverage was in accordance with government reluctance to rapidly increase the financial burden on firms. In accordance with this line of reasoning is that the social insurance law, passed in October 2010, states not the ambition of full coverage, but rather that the "social insurance level shall match with economic and social development level".

I focus on the role of the "tightness" of the labor market for firm provision of UI. This concept of tightness denotes the number of unemployed workers available on the labor market relative to the number of other firms that are also looking to fill vacant spots on their payroll. I argue that firms which are looking for labor on tighter labor markets have increased incentives to provide unemployment insurance because doing so increases the firm's attractiveness and makes it easier to find new workers.

The main finding of the essay is that the labor market situation does affect firm behavior. On tighter labor markets, firms are more likely to offer UI, but the effect is not large. In comparison to the situation on the labor market, the empirical analysis suggests other factors that may be important determinants of firm behavior. Most notably, firm ownership, its cost profile, and characteristics of its workforce emerge as potential avenues for future research.

The variation in insurance provision across ownership sectors is remarkable. On average, I find that firms with foreign ownership are twenty percent more likely to provide insurance compared to private domestic firms.[9] Considering this result through the lens of globalization, it indicates that increased foreign ownership can be positive for worker welfare. When judging firms competitiveness on the Chinese market, it could instead indicate an uneven playing field to the disadvantage of foreign firms, in particular when considering the high total cost of insurance provision.[10]

Finally, there are indications that the firm's cost profile is important. The results show that insurance provision is more likely if the firm has a large market share, if it receives state subsidies, or if it is located in a place with a lower total cost-level for social insurance. Moreover, unionized firms and firms with more highly educated workforces are more likely to be insurance providers.

Institutions, natural resources and macroeconomic stability

When comparing transition economies, one aspect that often stands out about the Chinese transition is its successful establishment of a globally competitive manufacturing sector. In contrast, Russia's transition led to a collapse of industrial production as GDP fell by approximately 30% between 1992 and 1998[11]. Regarding this difference, an often mentioned reason that China and Russia cannot easily be compared is the Russian economy's re-

[9] After controlling for various measures of firm location, ability to pay, cost profile, workforce characteristics, and the local institutional setting.

[10] Total costs for the main benefit programs: unemployment insurance, pensions, health care and housing, together exceed 50% of the total unit labor cost (Banister, 2005).

[11] Author's own calculations based on IMF's World Economic Outlook Database, available online at www.imf.org.

liance on natural resource exports. In **essay 4**, I examine the effect of the value of such exports on a country's macroeconomic stability, focusing on oil-exporters as a case study.

The essay takes as a starting point the controversy surrounding the role of natural resource prices in a country's external competitiveness, as measured by its real exchange rate.[12] In particular, for countries that export oil, the price of oil has in some cases been shown to substantially impact on the price of the country's export goods on the global market, and in some cases not. Investigating this controversy is important because of the rapid price movements that characterize natural resources, and where the oil price is a case-in-point. A situation in which the real exchange rate of a country follows the movements of a volatile resource price is detrimental to growth (e.g. Aghion et al., 2009).

In the essay, I propose, and empirically test, the hypothesis that a country's political and legal institutions determine how its real exchange rate is affected by the oil price. Institutions are likely to mediate the price effect because they determine how careful (or not) governments are in spending their natural resource wealth. As observed in the extensive case studies of oil-exporting economies conducted by Gelb (1988), the most common characteristic of these countries' fiscal behavior was the systematic overspending of resource revenues.

I use data for 33 oil-exporting countries in 1985-2005 to test the hypothesis that institutions affect the relationship between the real exchange rate and the oil price. I find that this is indeed the case, but not all institutions are equally important. Strong bureaucracies and legal systems are found to be most relevant. This observation echoes the argument from political economy models of "rent-seeking" (Tornell and Lane, 1999; Mehlum, 2006). The important institutions to keeping revenue spending responsible appear to be the ones that reduce the incentives for interest groups to engage in rent-seeking, rather than independent and productive, activities.

The importance of the essay's result is demonstrated by listing the oil-exporting countries in the dataset over the institutional dimensions shown to be important for insulating currencies from the oil price. The countries that top the list in terms of institutional quality, Norway and Saudi Arabia, are the same countries for which the macroeconomic literature has observed a "puzzling" absence of a strong role of the oil price in determining real exchange rates.

[12] The real exchange rate is a measure of how the price of a country's exports compares to the price of the country's imports.

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Essay 1: Income Inequality between Chinese Regions: Newfound Harmony or Continued Discord?*

1. Introduction

China is the most populous country and has the fastest growing economy in the world. It has maintained an annual average growth rate of 10% over the past 30 years, and surpassed Japan to become the world's second largest economy in 2010. The economic reform agenda that spurred this growth is highly diversified geographically. With the explicit goal of letting some provinces "get rich first", preferential policies aimed at first attracting technology and capital to localities along the coast. However, rather than the intended trickle down of growth to inland provinces, income inequalities between and within regions arose rapidly (Fan et al., 2009). Re-balancing these inequalities and achieving balanced, or "harmonious", growth continues to be a priority of the Chinese government.

A key prediction of the neoclassical growth model (Solow, 1956) is that of economic convergence, meaning that for given endowments, poorer economies will eventually catch up with richer ones. This prediction has generated a large amount of empirical research on the existence of convergence between countries, and between regions within countries. In general, studies using cross-sectional methods have found evidence of conditional convergence (e.g. Baumol, 1986; Dorwick and Nguyen, 1989; Barro, 1991; Barro and Sala-i-Martin, 1992; Mankiw et al., 1992) while time series studies have concluded that, according to the tests applied, economies do not converge (e.g. Bernard, 1992; Bernard and Durlauf, 1995; Quah, 1993).

One possible reason for the discrepancy in the empirical literature is that time series studies tend to use an inflexible test of convergence, requiring the difference between the two output series to be stationary with a zero mean[1]. As noted, for example by Harvey and Bates (2003), Carvalho and Harvey (2005) and in passing by Bernard and Durlauf (1996), this approach actually

* This paper is co-authored with Johan Lyhagen, Uppsala University. The authors gratefully acknowledge helpful comments from Tor Eriksson, Bertil Holmlund, and Joakim Westerlund.
[1]Following Proposition 5 of Bernard and Durlauf (1996), or equivalently in a panel setting as proposed by Evans (1998).

means testing whether convergence has already happened, which is completely different from the question of whether convergence is occurring.

In this paper we develop a new time series test that can determine whether income levels are in the process of becoming more similar or not, thus getting closer to testing the true convergence hypothesis. The test is constructed by introducing a nonlinear trend into the vector error correction model. By specifying this nonlinear trend as a logistic function we are able to estimate the trajectory of convergence or divergence by extracting the parameters associated with the estimated trend[2]. We can thus answer a number of hypotheses relating to the differences in growth pattern between China's regions[3]. In particular, we can plot the estimated trends for any cointegrating relationships between regions, and use the plots to determine graphically whether regions have: i) converged to the same income level (absolute convergence) or diverged to different income levels, ii) converged to different, but stable, income levels (conditional convergence), or iii) diverged and then converged, or vice-versa. In addition, we can analyze the timing, persistence, and changes between these situations over time.

The test is applied to Chinese regional data for the period 1952-1977[4]. We first consider the question of nationwide convergence, and then test for convergence within two classifications of regional subgroups. First, we divide regions into either Coastal or Central areas, the major political divide in the market reform period, during which the Coastal regions benefitted greatly from the Open Door policy. Next, we examine the existence of convergence among the regional sub-divisions of the North East, South West, North West and Central areas. Here, the North East represents the provinces that were the main locations of the State-run heavy industries during the planned economy period, while the Central and South Western areas are cross-sections covering leading Coastal regions and their hinterlands. Finally, the North West area includes poor interior regions that lack coastal access.

The economic performances of China's 32 regions represent a suitable test bed for our model. Many of the regions are the same size as entire countries, but share institutions, language, and cultural traits which could cloud convergence tests in a cross-country setting. Previous empirical analyses of the convergence question also mirror the methodological divide mentioned above: cross-sectional studies have provided evidence in favor of economic convergence (Jian et al., 1996; Chen and Fleisher, 1996; Raiser, 1998;

[2] A similar approach is the non-linear panel unit root test used by Lau (2010). There are some drawbacks with that method, e.g. the theory applied postulates that all differences between pairs of regions should cointegrate and that the test used only has the alternative hypothesis that a proportion of the differences is stationary.

[3] Following the literature, the term "region" is used to denote China's 22 provinces, 4 self-governing municipalities, and 5 Autonomous Regions. These three types of entities operate at the same administrative level and are directly subordinate to the central government.

[4] We also considered the reform period, 1978-2007, with very similar results to the ones presented for the longer period.

Weeks and Yao, 2003). In contrast, time series studies have documented economic divergence (Pedroni and Yao, 2006; Westerlund et al., 2010). Using a more flexible convergence test has the potential for shedding light on the roots of this controversy.

Our paper does not simply contribute to the theoretical literature associated with time series tests of economic convergence; it also deepens our understanding of China's transition experience with respect to the critical question of income inequality. Our results show that divergence of economic output between roughly half of the region pairs started with market reform in the late 1970s, and showed no sign of reversal by 2007. Simultaneously, the other half of the region pairs grew at varying speeds while their income difference remained relatively constant; this can be interpreted as indicating that conditional convergence was achieved.

The paper is organized as follows. Section 2 provides a literature review of studies of Chinese regional income convergence. Section 3 presents the data and Section 4 the model. Empirical results are contained in Section 5, and Section 6 outlines our conclusion.

2. Literature review

Early studies of economic convergence used cross sectional and panel data to produce regressions of regional growth rates on initial per capita income levels. These studies employed the term "unconditional convergence", also called β-convergence, to describe a negative coefficient on initial income (Barro and Sala-i-Martin, 1997). Conditional convergence, or σ-convergence, in contrast, denotes the existence of a negative relationship after the inclusion of a set of regional covariates that are expected to proxy for permanent differences in regional steady state income levels (Barro and Sala-i Martin, 1997).

In the Chinese case, cross-section and panel data studies that have used parametric regression methods have found evidence of economic catch-up of poor regions with richer ones. Jian, et al. (1996) did not find convergence in the pre-reform period, but showed that absolute convergence emerged strongly as a result of economic reform after 1978. Chen and Fleisher (1996), and Li et al. (1998), found evidence to support both absolute and conditional convergence in the period 1978-1993. Cai et al. (2002) and Raiser (1998) added support for conditional convergence as a consequence of market reform, and more recently, Zou and Zhou (2007) documented convergence within one group "developing" and one group of "developed" provinces, with faster convergence noted in the latter

Weeks and Yao (2003) note that the cross-sectional methodology to measure convergence can suffer from bias if the variables included fail to account for: i) unobserved province-specific heterogeneity in initial technol-

ogy levels and ii) endogenous explanatory variables. Using Generalized Method of Moments estimation with province- and time-specific effects, they found that provinces were diverging in the pre-reform period, but that the interior and coastal regions converged to parallel underlying steady-states after 1979.

Following criticism of the parametric regression methods used to investigate the convergence hypothesis (Quah, 1993, 1997; Maasoumi and Wang, 2007)[5], researchers increasingly turned to time series analysis. Noting the common presence of unit roots in income per capita time series, Bernard and Durlauf (1991) and Quah (1993) proposed that stationarity of the difference between the income time series of two regions could be interpreted as convergence. The criterion was modified by Evans (1998) to fit panel data.

Pedroni and Yao (2006) employ Evan's method to investigate nationwide convergence between Chinese regions as well as convergence within various subgroups of regions in the pre-reform and reform periods. The authors construct a panel of relative income series by subtracting the average real income of the group under consideration (the cross-sectional mean) from the real income of each group member at each time point. Following Evans (1998), stationarity of the panel is interpreted as convergence within the group. Contrary to the results presented in earlier studies, the authors find evidence of income divergence, both nationally and within the vast majority of the examined subgroups. Lau (2010) extends the analysis of Pedroni and Yao by allowing for a non-linear unit root test, concluding that the main conclusions remain the same.

Westerlund et al. (2010) note that Evans' (1998) method, and the conclusions drawn by Pedroni and Yao (2006), may be sensitive to the choice of provinces included in each panel (and thereby in the calculation of the cross-sectional mean). Rather than relying on this calculation, they therefore conduct joint tests for unit roots in the income differences time series of all possible region pairs. The result, however, is no different: regions diverge economically, both nationally and in various sub-samples (Westerlund et al., 2010).

[5] Evans and Karras (1996) argue that the conventional cross-sectional technique of testing for economic convergence produces invalid inferences unless all permanent cross-economy differences in per capita GDP are perfectly controlled for and unless the countries or regions under consideration have income series that exhibit identical first-order autoregressive dynamic structures.

3. Data

The Hsueh and Li (1999) GDP data for Chinese regions in the period 1952-1995 are combined with data from China's Statistical Yearbook for 1996-2007.[6] Regional GDP per capita series at 1995 prices are calculated using information on total regional population and regional GDP deflators from China's statistical yearbook. As in previous studies, we exclude the regions of Hainan and Tibet, the two smallest regional economies, for which data are missing prior to 1987 and 1988, respectively. To allow a comparison over time, Chongqing is treated as part of the Sichuan region from which it was separated in 1997. Information on the geographical classification of our regions is presented in Table 1. We use two categorizations: the division "Interior" vs. "Coastal", and the division between "Central", "North East", "North West", and "South West" areas.

Figures 1-3 provide a graphical overview of the general trends with respect to income convergence. Using first the classification of regions as belonging to either the Interior (10 regions) or the Coastal area (18 regions), Figure 1 shows the mean income level per capita of the two areas relative to the national average. We see that the average Coastal region developed faster than the average Interior region, and that the income gap between these two areas widened during the reform period. Figures 2 and 3 illustrate the within-group heterogeneity by plotting the growth paths of the poorest and wealthiest two regions within the Coastal and Interior areas, respectively. Figure 2 shows that the richest Coastal province (Shanghai) grew at a roughly similar pace to the poorest Coastal region (Hebei). Figure 3 shows a similar pattern for the regions located in the Interior. During the sample period, the log of real GDP per capita in the poorest Interior region (Guizhou) remained at an apparently constant ratio to the wealthiest region (Beijing).

[6] Following Pedroni and Yao (2006); Lau (2010) and Westerlund et al. (2010).

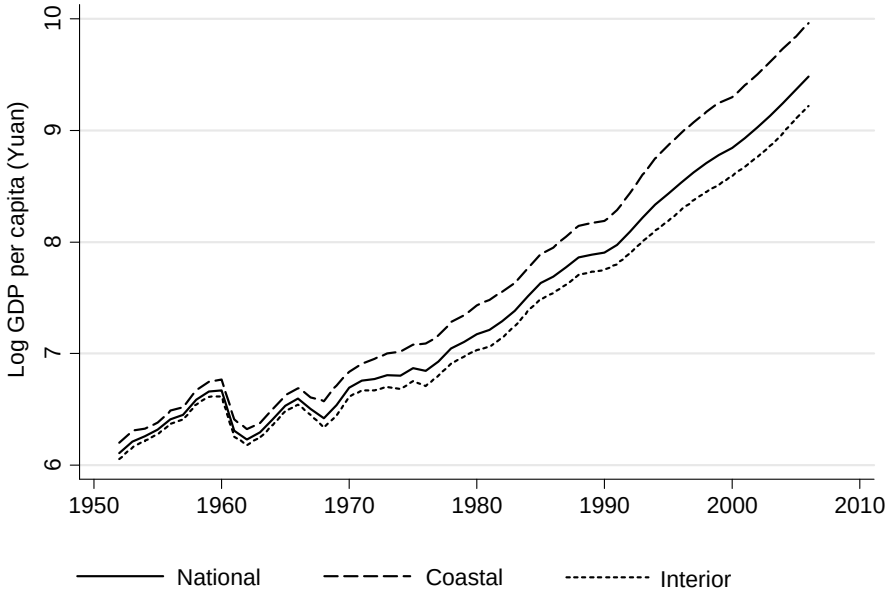

Figure 1: Average log GDP per capita for the full sample (National), and the Interior and Coastal areas, 1952-2007.

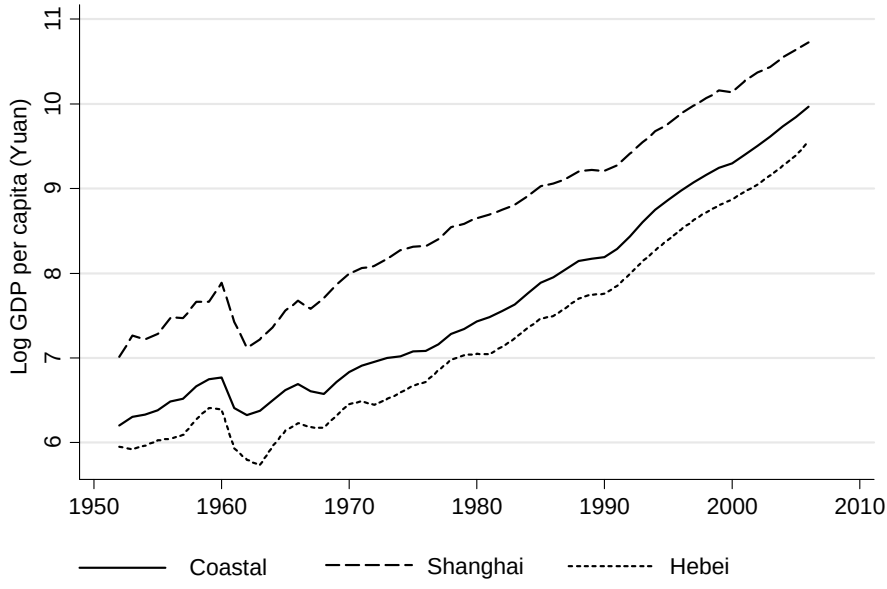

Figure 2: Growth paths for the poorest and richest Coastal regions relative to the average, 1952-2007.

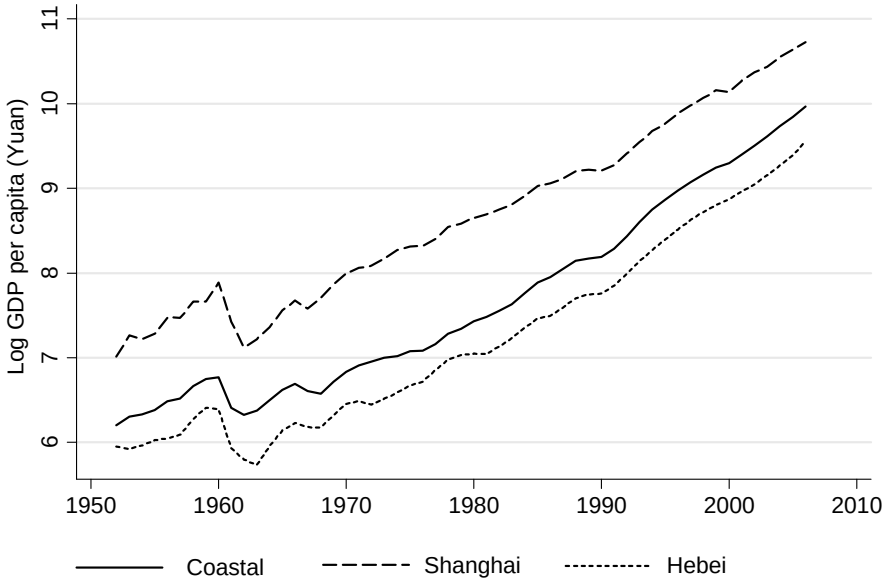

Figure 3: Growth paths for the poorest and richest Interior regions relative to the average, 1952-2007.

Table 1: Classification of regions into geographical areas.

Region	Classification 1	Classification 2
Beijing	Interior	Central
Tianjing	Coastal	Central
Hebei	Coastal	Central
Shanxi	Interior	Central
Mongolia	Interior	NW
Liaoning	Coastal	NE
Jilin	Interior	NE
Heilongjiang	Interior	NE
Shanghai	Coastal	Central
Jiangsu	Coastal	Central
Zhejiang	Coastal	Central
Anhui	Interior	Central
Fujian	Coastal	Central
Jiangxi	Interior	Central
Shandong	Coastal	Central
Henan	Interior	Central
Hubei	Interior	Central
Hunan	Interior	Central
Guangdong	Coastal	Central
Guangxi	Coastal	SW
Sichuan	Interior	SW
Guizhou	Interior	SW
Yunnan	Interior	SW
Shaanxi	Interior	NW
Gansu	Interior	NW
Qinghai	Interior	NW
Ningxia	Interior	NW
Xinjiang	Interior	NW

4. Model

This section describes the econometric method used in the paper. The data are transformed using natural logarithms. As there is a rather large number of regions, i.e. 28, it is not possible to conduct a full system/multivariate analysis.

As is well known, see references above, differences between pairs of regions or between a region and the average of all regions should be stationary although the GDP for each region is a unit root process. As pointed out by

Westerlund, et al. (2010), there are some problems with the analysis of pairs of regions and the average of the regions. The unit root approach implies that the cointegrating vector should be [1,-1] between each pair of regions. This is not necessarily true and the unit root test would, in this case, not reject the null hypothesis of a unit root.

Our starting point is the (bivariate) vector autoregressive model, VAR. The VAR, in levels, is used to choose lag lengths for each pair of regions based on the Schwarz information criterion. The next step is to perform the Johansen likelihood ratio test for the cointegrating rank. Because the growth of regions is interdependent, we conduct residual bootstrapping, in which the data are generated for the pairs using residuals from the same point in time. It is well known that the bootstrap method works considerably better than using asymptotic critical values, see e.g. Swensen (2006).

The problem with the cointegration approach described above is that it is only valid if convergence has already taken place, and not when convergence is occurring, as previously noted by Harvey and Bates (2003) and Carvalho and Harvey (2005). To deal with this we introduce a deterministic variable to capture this phenomenon. This variable is a logistic function[7]:

$$g_t = \frac{1}{1+e^{-\gamma(t-\delta)}}$$

The VECM becomes:

$$\Delta X_t = \mu + \alpha(\beta' \quad \rho')\begin{pmatrix} X_{t-1} \\ g_t \end{pmatrix} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t$$

where μ is an unrestricted constant, ε_t is a random vector with covariance matrix Ω etc., see Johansen (1995). It is possible to use this model to test for the number of cointegrating vectors; the asymptotic distribution of the likelihood ratio test is given by Johansen and Nielsen (1993). In addition, in this case, we bootstrap to obtain an approximation of the small sample distribution. Furthermore, it is possible to test the null hypothesis that $\rho = 0$ using the likelihood ratio test, although the distribution is non-trivial due to the non-identifiability of γ and δ under the nullhypothesis. The bootstrap approach is also used here. To determine whether there was first a period of divergence, and then a period of convergence (or vice-versa), a second logistic function is added and the tests for this situation followed those outlined above.

[7] Due to identification of the parameters, without loss of generality, the first observation is normalized to zero and the last to one.

The estimation of the non-linear VECM is non-trivial as it is non-linear with respect to the parameters, but, depending on γ and δ, the model is linear for the remaining parameters. Hence, given γ and δ it is trivial to calculate the likelihood and we use a numerical optimization routine to optimize γ and δ. This is performed using MATLAB version 2009a.

5. Empirical results

5.1 Cointegration tests

Table 2 presents the test results for the cointegrating rank of each geographical area under consideration, as well as tests of non-linearity in the trends of the cointegrating relationships. Column 1 presents the p-values for the standard Johansen test for no cointegration, $r = 0$, when including a linear trend and a constant in the cointegrating relationship. The null hypothesis is rejected for all samples except the North Eastern and South Western subgroups. These rejections can be interpreted in two ways. On one hand, it is more difficult to reject the null hypothesis of no cointegration when we have fewer observations. On the other hand, the test may erroneously show no cointegration if the cointegrating relationship has a non-linear rather than a linear trend.

Next, we conduct the Johansen test for the null of no cointegration $r = 0$ versus of a cointegrating relationship with either 1) a nonlinear trend with one nonlinear element (Column 2) or 2) with two nonlinear elements (Column 3). The results clearly indicate that we should reject the null hypothesis for all sub-samples in the cases of both trend specifications. Thus, the results support the previous conjecture regarding the erroneous acceptance of the null hypothesis (i.e. no cointegration) in Column 1 due to incorrect specification of the trend. We could however still have a situation of no cointegration if the rank is two rather than one, in which case both time series constitute separate stationary processes. Testing for this property (Column 4) does not lead us to reject the null hypothesis of one cointegrating relationship for any of the samples.

Having reached the conclusion that there is one cointegrating relationship, we next investigate the statistical significance of the nonlinear trend parameters. The results presented in Columns 5-6 relate to applying the likelihood ratio test to a single non-linear element (Column 5), or to two nonlinear elements (Column 6). The p-values in Table 2 indicate strong support for the statistical significance of the trend parameters, regardless of the test.[8]

[8] A test based on the Taylor approximation principle presented by Teräsvirta (1994) was also carried out. The result was in accordance with the other tests, hence not reported.

We still need to exclude the possibility that there are two stationary processes (no cointegration) and nonlinear trends. Column 7, therefore, shows the result of testing the hypothesis that there is one cointegrating relationship $(r = 1)$ against the alternative that there are two cointegrating relationships $(r = 2)$ with nonlinear trends. In this case, we consistently accept the hypothesis that there is one cointegrating relationship.

The last step of the analysis is to determine the precise nonlinear characteristics of the trends. The final column in Table 2 shows the results of comparing our two trend specifications, where the null hypothesis is specified as the existence of one nonlinear element. The p-values indicate that we should accept the null hypothesis in half of the cases. It is accepted for the Coastal and the Central areas. For the Interior area, the North West, the North East, and the South West we instead find evidence of double nonlinearity.

Table 2: P-values for the hypotheses relating to cointegrating rank and functional form of the deterministic term for various groups, 1952-2007.

Hypothesis/group	1	2	3	4	5	6	7	8
All	0.008	0.000	0.000	0.120	0.000	0.000	1.000	0.000
Coastal	0.055	0.000	0.000	0.070	0.000	0.000	1.000	0.000
Interior	0.003	0.000	0.000	0.170	0.000	0.000	1.000	0.261
Central	0.025	0.000	0.000	0.138	0.000	0.000	1.000	0.000
NW	0.040	0.000	0.000	0.411	0.000	0.000	1.000	0.985
NE	0.160	0.008	0.038	0.218	0.000	0.003	1.000	1.000
SW	0.306	0.003	0.045	0.175	0.000	0.000	1.000	0.223

Notes: The hypotheses were as follows, where the type of deterministic term is restricted to the long run relationship and r is the cointegrating rank: 1) constant, $H_0 : r = 0, H_A : r = 2$; 2) constant and logistic function, $H_0 : r = 0, H_A : r = 2$; 3) constant and two logistic functions, $H_0 : r = 0, H_A : r = 2$; 4) constant $H_0 : r = \leq 1, H_A : r = 2$; 5) constant and $r = 1$, H_0 : linear, H_A : two logistic functions; 6) constant and $r = 1$, H_0 : linear, H_A : two logistic functions; 7) constant and logistic function $H_0 : r = 1, H_A : r = 2$, 8) $r = 1$, H_0 : one logistic function, H_A : two logistic functions.

5.2 Graphical analysis of nonlinear cointegrating relationships

It is important to remember that providing empirical evidence for the existence of a nonlinear trend in the cointegrating relationship between the real GDP per capita series of two regions does not mean that we have evidence of convergence. If the trend is upwards, this means that the regions are in

fact diverging. Only a negative trend shows that the income levels of two regions are becoming more similar.
similar.

The main conclusions about convergence or divergence between China's regions are drawn by examining graphical plots of the cointegrating relationships between the regional pairs. By studying these plots, we are able to address three key issues. First, we want to examine the slope of the nonlinear trends to determine whether the region pairs experienced divergence (upward slope) or convergence (downward slope). Second, we want to evaluate the magnitude of the gradients of the slopes. Steep lines indicate rapid processes of convergence or divergence. Straight lines indicate that the region pairs have grown while keeping a constant income gap. Thus, we can interpret straight lines as conditional convergence in the case of a nonzero mean, and unconditional convergence in the case of a zero mean. Third, we want to inspect visually the distribution of trends in each group of regions to determine the relative occurrence of convergence or divergence.

Using the information in Column 8 of Table 2 to guide the choice of trend specification, we plot region pair trends using either single or double nonlinearity. Because of the large number of region pairs, for each given year the trends are first ranked by size, and the distribution of trends then illustrated by quantiles. Looking at Figure 4, which illustrates the distribution of region-pair trends for the nationwide sample, we can interpret the lines in the figures as follows. The middle of the five indicator lines shows the median trend, meaning that 50% of all the region-pair trends lie above this line. The second indicator line from the top delineates the top 25% of the trends, and 0.025% of the trends are above the uppermost line. Analogously, the two bottom lines indicate the lowest 25% and 0.025% of the trends.

In Figure 4, the distribution of trends shows a clear pattern of upward sloping nonlinear trends for more than 50% of the regional pairs. Focusing on the year indicators on the X-axis, we can also observe that divergence, as indicated by the upward sloping trends, started during the 1970s, but accelerated for some region pairs from the early 1980s. Focusing on the right side of the graph, there are no directional changes in the trends. This demonstrates that, for the diverging 50% of the regions, the process of divergence has not lost pace in recent years. For the remaining and non-diverging pairs, we can see flat trends, most with nonzero means. Thus, the results do not support ongoing convergence, but rather that the regions were growing along separate but different trajectories, retaining constant income differences rather than becoming more similar.

We next dissect the nationwide pattern and examine smaller groups of regions to see whether we can identify areas of divergence and conditional convergence. Figure 5 shows the trend distributions of the region pairs in the Interior and the Coastal areas, respectively. Comparing the two figures, we note that the trend distributions appear very similar, with a mixed picture of

divergence and convergence between the various region pairs. Upward trends show that divergence has occurred between more than 50% of the pairs in each of the two areas. However, there are also slightly steeper gradients, that is, faster divergence, for the most rapidly diverging pairs in the interior area. With regards to the timing of divergence, the coastal regions appear to have grown apart more consistently over the entire 1952-2007 period. In contrast, we note that divergence in the interior regions appears to be more directly associated with the economic reforms of the past three decades.

Comparing the regions in the lower part of Figure 5, we see that many of the flatter trends for the cointegrating relationships among the Coastal regions are associated with means that are closer to zero. This implies that the pairs of Coastal regions that have grown at very similar rates have had more equal income levels, while there are larger and persistent income differences among the regions in the Interior part of the country.

Smaller subdivisions of the region pairs are presented in Figure 6, which shows the distribution of trends for the region pairs within the North East and South Western groups, and Figure 7, which shows the trends for the region pairs in the Central and North Western regions.

In the North Eastern provinces of Heilongjiang, Jilin and Liaoning the figure indicates that the income levels for these regions have not converged. Rather, the pace of divergence accelerated in the post-1979 period, indicating that the dramatic privatization and retrenchment of state-owned industries was conducive to increased growth dispersion between these three regions. In the South Western regions of Guangxi, Sichuan, Giuzhou and Yunnan the growth pattern appears more stable than for the Northern coast. Only one province pair diverged markedly in the post-1979 period.

Finally, we consider the Central and North Western areas. Because of the larger number of regions in the Central sub-division, the figure shows the trend distribution rather than the individual region pair trends as for the North West. In these regions, the growth pattern in both areas mirrors the nationwide results. While about half of the regions have followed diverging growth trajectories, there is also a substantial proportion of regions which have neither diverged nor converged, but rather continued to grow whilst their income differences have remained constant. In some cases, these differences are very small, indicating that many regions in fact experienced quite similar income levels during the entire period under investigation.

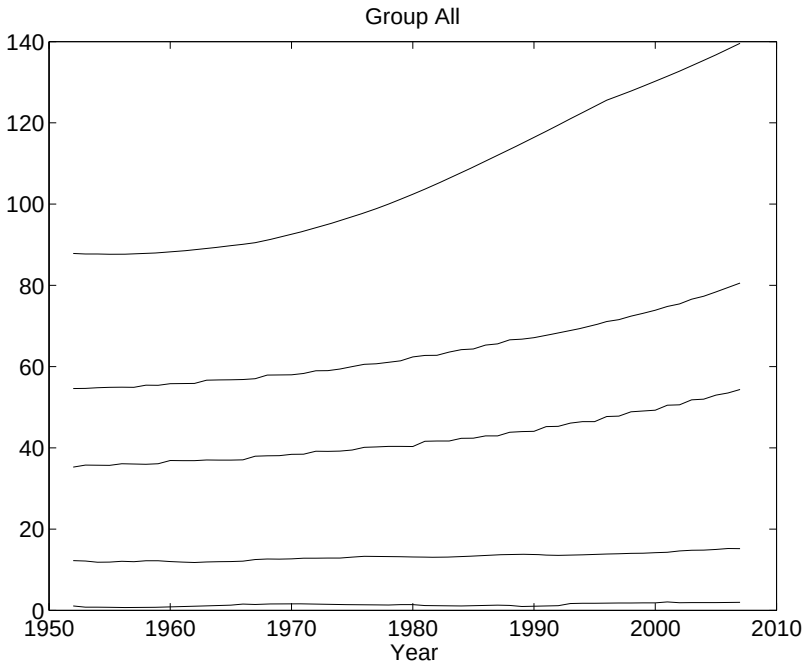

Figure 4: Quantiles 0.025, 0.25, 0.5, 0.75, and 0.975 of the absolute value of the long run deterministic terms for all region pairs.

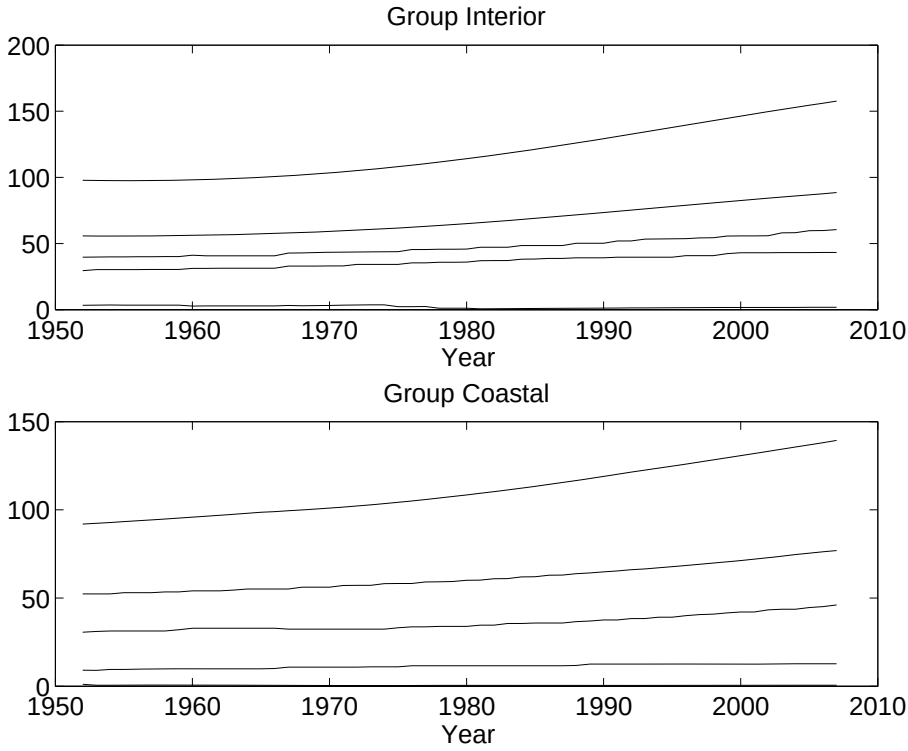

Figure 5: Quantiles 0.025, 0.25, 0.5, 0.75, and 0.975 of the absolute value of the long run deterministic terms for all region pairs in the Coastal and Interior geographical groups.

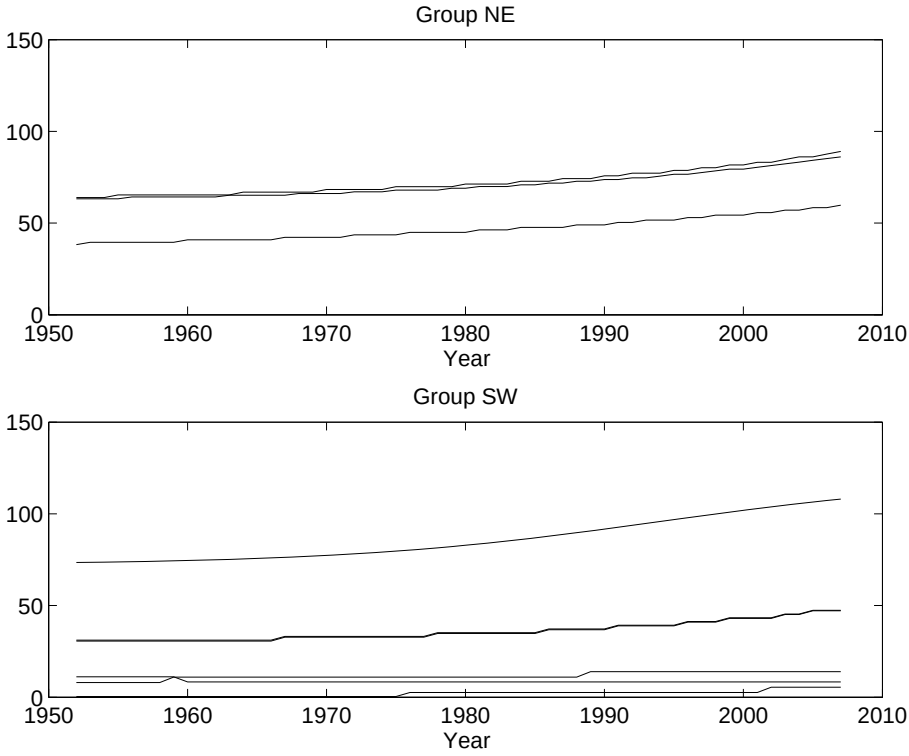

Figure 6: Absolute values of the individual deterministic trends for the North East and South West groups.

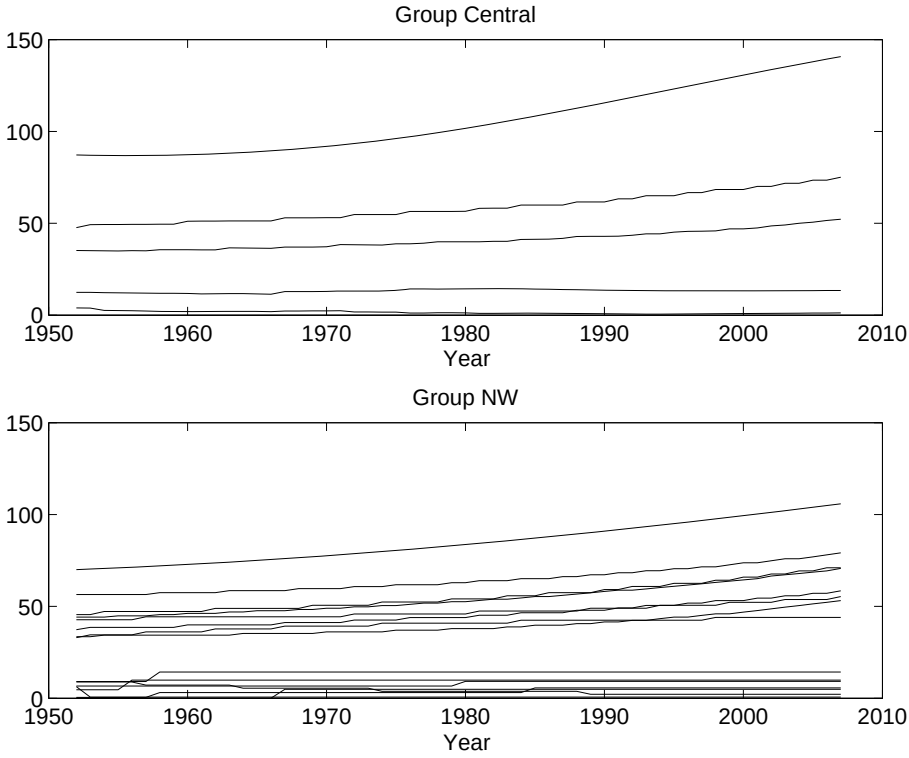

Figure 7: Quantiles 0.025, 0.25, 0.5, 0.75, and 0.975 of the absolute value of the long run deterministic terms for all region pairs in the Central geographical group, and the absolute value of the individual deterministic trends for the North Western group.

6. Conclusions

This paper presents a new methodology to test for economic convergence or divergence. Using Chinese regions as a case study, we found that the test provides useful insights into the country's internal growth patterns. Cointegration tests within the VAR model led to the rejection of the hypothesis of cointegrating relationships with linear trends in favor of the occurrence of nonlinear trends. This demonstrates that our test specification captures more of the growth dynamics than previously suggested methods (e.g. Evans, 1998). Hence, our approach is more likely to produce correct answers to the convergence question.

Besides the proposed model's better fit to the data, the ability to illustrate the trends associated with the cointegrating relationships also yielded a number of insights. Unlike previous time series studies, we were able to show that economic divergence accelerated in the reform period without reducing the number of observations by dividing the data into pre- and post-reform samples. Moreover, in allowing for both single and double nonlinearities, we were able to reject the hypothesis that regions had diverged at the beginning of the reform period and then converged in more recent years. Thus, there is little evidence of success for the last decade's investment efforts aimed at bridging the economic gap between the East and West by boosting the development of the Western regions.

Compared to previous time series studies of convergence in China, our method allowed a flexible assessment of convergence and divergence nationwide and within groups of regions, rather than the need to conclude in favor of either one pattern or the other (e.g. Pedroni and Yao, 2006; Westerlund et al., 2010). Examining the trends of the cointegrating relationships between regional pairs, we found that although divergence was taking place among approximately half of the pairs, the other half neither diverged nor converged during the period 1952-2007. Instead, the income differences between these regions remained roughly constant, suggesting that there was conditional convergence.

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Essay 2: Gender and Work Compensation in China's Industrial Sector*

1. Introduction

An investigation of gender-related earnings discrimination in China is urgently required. Not only is the question of great economic and social importance to the female half of the labor force, which reached 640 million in 2004, it is also of wider economic importance as discrimination discourages groups of workers from participating in employment and investing in their own education and skill development. This under development and underutilization of human capital damages long-term economic growth.

In this paper, I assess the gender wage gap in China's urban industries by drawing on a large and nationally representative firm-level dataset. The main objective is to examine how women with different levels of education fare within their industrial workplaces. By testing whether women's average earnings correspond to their contributions to productivity, I examine if they are subject to wage discrimination. The analysis is then expanded to investigate the role of ownership, industry, and region on women's relative wages and productivities.

The paper addresses a number of limitations in the existing literature. First, the vast majority of previous empirical studies of gender-based wage discrimination in China uses data on the individual or household level, and interpret a residual wage gap as being indicative of discrimination (see Shen and Deng, 2008 for a survey). But, as pointed out by e.g. Altonji and Blank (1999), the observable individual-level characteristics in the wage-equation are unlikely to capture all gender-related differences in productivity. This method therefore risks overestimating the degree of discrimination[9].

By employing the method proposed by Hellerstein and Neumark (1999) and Hellerstein et al. (1999), I use firm-level data to obtain direct measures

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9 The part of the wage-gap not explained by proxies for individual productivity is indeed large in previous studies, for example 53-63% found by Gustafsson and Li (2000), 89-92% by Liu et al (2000) and 85% by Wang and Cai (2008).

of the productivities of men and women from the production function[10]. These productivities are then compared to the average marginal wages of the same groups of workers. To maximize profits, firms should hire the most productive workers for the lowest costs, so that wages for all worker groups should match their productivities. If, instead, we find deviations between wages and productivity, this observation is consistent with employers sacrificing profits in order to satisfy a preference for particular groups of workers (Becker, 1971). In the case of a transition country, sub-optimal worker selection could also be a legacy from the planned economy. In this period, there was a misallocation of workers to jobs, and wages were not set to reflect productivity, but in accordance with a fixed wage scheme (as argued by Dong and Zhang, 2009). Another possibility is that undeveloped capital and labor markets prevent firms from attaining their profit-maximizing size (Fleisher et al., 2004; in press).

A second contribution of the paper stems from the quality of the dataset. In this respect, the paper provides substantial improvements on Dong and Zhang's (2009) unique -level study of gender-related wage discrimination in China, which uses a methodology similar to the one employed in this paper. Compared to Dong and Zhang's study, the data coverage is expanded from a fraction of the economy – 856 firms located in a few large cities – to a sample of 255,568 firms. These data include all large and medium-sized industrial enterprises, as well as the majority of those firms designated as being "small", thus reducing the risk of selection bias[11] and increasing statistical precision. Moreover, gender-disaggregated information on the education levels of employees allows for the computation of skill-weighted wage- and productivity gaps between men and women.

The study is also the first to examine the role of non-wage compensation with respect to the gender-polarization of total earnings in Chinese industry. Providing employees with the main social insurances has been shown to carry a cost that approaches half of the total wage bill (Banister, 2005)[12]. Moreover, previous research has argued that insurance cover has become increasingly gender-biased during the transition period (Razavi, 2007; Wang, 2006).

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10 Using this popular methodology has the advantage of making the results comparable to previous studies examining China and other countries. These studies include, but are not limited to, analysis of the US (Hellerstein et al., 1999 and Hellerstein and Neumark, 2005); Israel (Hellerstein and Neumark, 1999); Germany (Bartolucci and Alberto, 2010); Japan (Asano and Kawaguchi, 2007); and China (Dong and Zhang, 2009).

11 This is particularly important in a vast economy such as that of China, where development levels, market reform intensity and social norms vary greatly across geographical regions.

12 Banister (2005) draws upon a survey conducted by China's Ministry of Labor, covering 11,704 urban enterprises in 51 large cities, to conclude that the standard wage measure of employee earnings should be increased by 53.8 percent to account fully for labor compensation actually paid by urban employers.

The results show an average skill-weighted gender wage gap of 12% in the nationwide sample. They do, however, not support the contention that social insurance payments contribute to further gender polarization of average earnings. At 12%, the average gender wage gap is smaller than those observed in other countries, based on similar methodologies and data[13].

Separating workers according to skill-level shows that unskilled women earn less than men, while skilled women earn more. This result corresponds to previous studies using individual-level data, which demonstrate a larger education premium for female than for male workers (Maurer-Fazio, 1999; Li, 2003). Further investigation of this result shows that it stems from wage setting in State-controlled sectors.

The highly representative dataset used here paints a picture of above-average proportions of female employees in light industrial sectors, in foreign-owned firms, and in firms located in the Coastal regions. In addition, women's average wage disadvantage is smaller in light industry than in other sectors. Thus, the results seem to indicate that women have gained both in terms of access to employment and in wage equality from the industrial policies associated with the market reform period.

With respect to gender wage discrimination, the results show that although women earn on average less than men, their relative productivity is even lower. For skilled women, this result holds for the State-controlled sectors, and for unskilled women it is true across all ownership, industry, and regional sub-samples considered. For the State-controlled sectors, these findings are interpreted as further evidence of a lingering mismatch between women and jobs as a result of State labor allocation. For the more recently established private sectors, an employer preference for unskilled female labor over male unskilled labor is confirmed by qualitative studies of Chinese factories. These studies highlight the preference for female unskilled labor motivated by women's perceived obedience and suitability for assembly line work (e.g. Pun, 2005).

The paper is organized as follows. Section 2 provides a brief overview of reforms to wage-setting and welfare systems in the post-1978 period and a review of previous research on gender-wage discrimination in China. Section 3 explains the basic econometric framework and section 4 discusses the dataset, variables and descriptive statistics. Benchmark estimation results and robustness checks are presented in section 5, and section 6 focuses on extending the dimensions of industry, ownership and location. Section 7 outlines my conclusions.

[13] It is also slightly narrower than the gap found by Dong and Zhang (2009), although their sample selection combined with the imprecision of the estimate makes a comparison difficult.

2. Wage and non-wage compensation in Chinese industry

In pre-reform socialist China, urban industrial workers enjoyed a system of guaranteed occupational and income security. This "Iron Rice Bowl" also provided access to social insurances (health care, pensions, education) that were distributed via state-owned work units. Equal rights on the labor market with respect to gender constituted an ethical commitment of the socialist state, and included juridical equality, the absence of gender bias with respect to entry into paid work, and access to social rights. Under this system, women's incomes improved radically: the gender-based wage differential narrowed and became small compared to that in other countries, particularly in urban areas (Croll, 1995). With an industrial policy that emphasized heavy and capital-intensive industry, many women occupied positions for which they were biologically disadvantaged compared to men, such as blue collar jobs requiring physical strength. However, the resulting skill mismatch was not reflected in the centrally determined and egalitarian remuneration system (Korzec, 1992).

In the late 1970s, a series of radical reforms were enacted with respect to the industrial labor market. The transformation of wage-setting started in the early 1980s when firms were given autonomy over their remuneration systems within government guidelines. By the 1990s, those guidelines had evolved into the imposition of minimum wages (Shen, 2007). Another centerpiece of reform was the transfer of the labor allocation decision from the state to individual enterprises. By 1994, firms had been given the right to dismiss workers. This new authority was extensively practiced during a massive retrenchment program in the state-owned sector. Between 1997 and 2002, more than 28 million state employees were asked to leave their jobs (Dong and Xu, 2009)[14].

Reform aimed at streamlining the organization of State-owned enterprises (SOEs) came as a response to increasing competition from the rapidly growing private and Foreign-financed sectors. These sectors also represented a problem from a social insurance perspective. A growing number of employees did not have access to the benefits provided by State-owned enterprises. Simultaneously, problems were brewing in those State-owned firms that were facing high insurance costs stemming from members of the workforce retiring and which had difficulties competing in the new economy. The main aim of the social system overhaul in the 1990s was to solve these two problems.

[14] Women were over-represented among the employees that were laid-off (Appleton et al., 2002; Giles et al., 2006) but simultaneously the demand for them increased in the export-oriented, and highly labor intensive, industries.

Put simply, social insurance reform consisted of breaking up the enterprise-based entitlement model, known as the Labor Insurance Scheme (LIS), into separate social insurances. Instead of the insurance provided by each firm individually under the LIS, the new programs aimed to share risk across firms by regional pooling of worker accounts. In this way, the 90s saw the birth of a two-tiered pension system in 1997, the Urban Employee's Basic Health Care Insurance System in 1998, an unemployment insurance scheme in 1999, and a system of workers' savings accounts for housing.

Although firms outside the state economy were obliged to join the new systems, compliance grew only modestly.[15] One reason was the high payment levels required to compensate for the pension arrears accumulated in State-owned firms.[16] Another important issue was the lack of a legal foundation for the social insurance programs, making enforcement difficult and evasion of obligatory participation an attractive option for firms.[17]

The social insurance requirements for rural firms and the rights to social insurance for rural workers are generally lower than those of their urban equivalents. The policies regarding rural workers (migrants) have varied widely between provinces and cities. Most localities have excluded migrants altogether, included them in an inferior rural system, or set up modified versions of their urban schemes. Some more ambitious localities, like Guangdong province, have brought migrants into the existing programs for urban workers (for further discussion, see Zhang et al., 2010).

Qualitative research has indicated an emerging pattern of gender-polarization in insurance cover, in particular in the export-oriented sectors of the economy (Razavi, 2007; Wang, 2006). It has also argued that women dismissed from SOEs were pushed into ownership sectors where social security protection was low (Cooke, 2001; Stockman, 1994). These claims have not, however, been verified quantitatively.

In contrast to non-wage compensation, there are a large number of studies dedicated to earnings discrimination with respect to wages. Most studies use some of the cross-sectional household datasets collected by the Chinese Academy of Social Sciences in 1988, 1995, and 2002 under a project known as CHIP (China Household Income Project). The common finding in these papers is a gender wage gap of slightly less than 20%, which was fairly constant over this period (Gustafsson and Li, 2000, Bishop et al. 2005, Démurg-

[15] In 2005, the pension scheme covered 17% of the urban workforce, and unemployment insurance covered 14% (NBS, National Bureau of Statistics of China, 2006, pp. 43 and 201). Overall health insurance coverage actually decreased between 1998 and 2003, as mainstream cover fell more sharply than the increase in commercial and other non-commercial insurances (Xu et al., 2007).

[16] Guidelines call for a total of 24% of the payroll to be directed to pensions accounts (Jackson et al, 2009), 16% to the housing accumulation funds (by 2003, Wang et al., 2005), 2-6% to medical insurance (Xu et al, 2007), and 2% to unemployment insurance (Vodopivec and Tong, 2008).

[17] The law was passed in October, 2010.

er et al., 2007, Shu and Bian, 2003). This result is corroborated by earlier findings from a Tianjin sample collected in 1993 (Bian et al., 2000), as well as the findings of researchers using data from the Urban Household Survey between 1988 and 1999 (Ng, 2004; Liu et al, 2004). Xu et al. (2006) report a slightly larger gap, 32%, in two recently urbanized towns in Zhejiang Province between 1999 and 2000. This is similar to Dong and Zhang's (2009) finding based on late 1990s firm-level data. With respect to the size of the wage-gap across educational categories, a number of studies report a smaller gender difference among highly educated employees (Xu et al., 2006; Hughes and Maurer-Fazio, 2002; Gustafsson and Li, 2000).

In previously published studies, the share of the gender-wage gap not explained by gender-related productivity proxies in the wage equation is generally large. Over 50% of the wage gap is left unexplained in some studies (Bishop et al. 2005, Shu and Bian, 2003, Gustafsson and Li, 2000). In others, it exceeds 75% (Liu et al. 2000; Rozelle et al. 2002; Wang and Cai, 2008). This evidence for substantial gender-based wage discrimination is contested by Dong and Zhang's (2009) study of firms. They show that women's estimated wages, although lower than men's, are not statistically different from their estimated productivity.

In a transition economy, the wage-setting practices of employers may differ widely between the large ownership sectors within the labor market. Xu et al. (2006) find the widest wage gap in privately owned enterprises, while Maurer-Fazio and Hughes (2002) and Maurer-Fazio et al. (1999) obtain the same result for firms registered as joint ventures. The studies by Liu et al. (2000) and Hughes and Maurer-Fazio (2002) add evidence that the widest wage-gaps exist in the most marketized ownership sectors. However, their results also show that these sectors have the largest proportion of the wage-gap that is not explained by observed worker characteristics. Rozelle et al. (2003) do not find any systematic association between the level of wage discrimination and the degree of market orientation by industry or ownership. Finally, the firm-level study by Dong and Zhang (2009) demonstrates that women are rewarded in accordance with their productivity in private firms but over-compensated compared to their productivity in the state-owned sector.

3. Analysis of earnings discrimination using firm-level data

Following Hellerstein and Neumark (1999), I use firm-level data to analyze the difference between women's and men's marginal wages and productivities. First, the production of firms, in value added terms, is expressed using a

Cobb Douglas production function in capital (K) and labor (L) [18]. The labor input is represented by a quality of labor index

$$QL = M_U + \phi_{FU} F_U + \phi_{MU} M_S + \phi_{FS} F_S \tag{1}$$

where F_j and M_j are the number of skilled, $j = S$ or unskilled, $j = U$, female or male employees. Hence, using natural logarithms, the full production function is expressed as

$$\begin{aligned} \ln(VA) &= \ln(A) + \alpha_K \ln(K) \\ &+ \alpha_{\cdot L} \ln[M_U + \phi_{FU} F_U + \phi_{MU} M_S + \phi_{FS} F_S] \end{aligned} \tag{2}$$

The specification of the quality of labor index implies that members of the four gender-skill groups are perfectly substitutable inputs, but that the marginal productivities of the groups may differ. The ϕ_{ij} parameters denote the average marginal productivity compared to male unskilled workers for the three other demographic groups: unskilled females, skilled females and skilled males. In this setting, $\phi_{FU} = \phi_{MU} = \phi_{FS} = 1$ would indicate that all four gender-skill groups contribute equally to the firm's output. Alternatively, $\phi_{FU} = 0.75$ would imply that the average marginal productivity of female unskilled employees is 75% that of the average for their male unskilled co-workers. Letting L represent total firm employment such that $M_U = L - F_U - M_S - F_S$ we can rewrite (2) as

$$\begin{aligned} \ln(VA) &= \ln(A) + \alpha_K \ln(K) \\ &+ \alpha_{\cdot L} \ln[L + (\phi_{FU} - 1)F_U + (\phi_{MS} - 1)M_S + (\phi_{FS} - 1)F_S] \end{aligned} \tag{3}$$

We next turn to the estimation of relative differentials for work compensation. A firm-level wage equation is set up in the form of a Mincer-type earnings equation

$$\begin{aligned} \ln(W_T) &= \lambda_0 + \ln(QL) = \lambda_0 \\ &+ \ln[L + (\lambda_{FU} - 1)F_U + (\lambda_{MS} - 1)M_S + (\lambda_{FS} - 1)F_S] \end{aligned} \tag{4}$$

[18] As discussed by Griliches and Ringstad (1971), the value added specification of the production function improves comparability of data across industries and across establishments within industries. It also allows greater comparability when industries or establishments differ in the degree of vertical integration, and can be derived from quite polar production function specifications: one in which the elasticity of substitution between materials and value added is infinite (i.e., Y=f(K,QL)+M); and one in which this elasticity of substitution is zero (so that materials have to be used in a fixed proportion to output).

where the dependent variable is the total wage bill, or the total amount of worker compensation, paid by the firm[19]. Analogous to the specification of the production function, the parameters λ_{ij} denote the relative average wages of the three gender-skill groups compared to the wages of unskilled male workers. A vector comprising control variables is included in both equations. It incorporates measures of geographical location, industry, size, age, ownership, unionization, and township-and-village enterprise status.

Whether worker groups are compensated in accordance with their productivity is tested by comparing the estimated results of the production function (3) and of the wage equation (4). In practice, this means that the size of the relative average wage parameters λ_{ij} are compared to the relative average productivity parameters ϕ_{ij} for each gender-skill group. Given the specification of the production function, is would be inconsistent with profit-maximizing or cost-minimizing behavior for firms to employ workers if their wage cost surpasses their productive value to the firm. Hence, such deviations could imply that firms sacrifice some profits to indulge in discriminatory preferences across worker groups (Becker, 1971). The deviations could also imply a legacy of misallocation of labor during the planned economy (Dong and Zhang, 2009), imperfect competition for labor, or that distortions within the product or labor markets have prevented firms from achieving their profit-maximizing size (Fleisher et al., in press; Hellerstein et al., 1999; 2005).

We are also interested in the wage- and productivity gaps between all women, skilled and unskilled, compared to all men. These are calculated by weighting the estimated wage and productivity gaps for the groups of skilled and unskilled workers by their proportions of the workforce in the sample of firms. Using the productivity gap as an example, the calculation becomes

$$\phi_F = \frac{\phi_{FU} P_{FU} + \phi_{FS} P_{FS}}{P_{MU} + \phi_{MS} P_{MS}} \tag{5}$$

[19] To see how this firm-level function can be understood as an aggregation of individual-level wage equations over workers in the firm, consider the individual level wage equation $w_i = w_M M_i + w_F F_i$ where w_i is the wage of an individual worker with gender *i*, w_M and w_F are average wages, and M_i and F_i are dummy variables for females and males respectively. Aggregating this function over the entire firm, the total wage bill is $W = w_M (L - F) + w_F F$, which can be expressed as $W = w_M [L + (\lambda_F - 1)F]$ where λ_F is the average relative wages of women compared to men, w_F / w_M. Taking logs gives a simplified equivalent of equation (4), in which the constant corresponds to the average wage of men $\lambda_0 = \ln w_M$.

where P_{ij} is the proportion of employees with gender *i* and skill level *j* in their respective gender group[20].

One limitation of the cross-sectional data is that we cannot distinguish between wage and productivity differences originating within or between firms. A major concern is that women may be sorted into lower paying and less productive firms, resulting in a downward bias in the estimated gender differences in wages and productivity. The latter part of this paper addresses this issue by re-estimating the model after separating firms into groups along the lines of three potential dimensions of sorting: by industry, ownership and location. Another robustness check for the total gender wage gap is achieved by adding firm data from 2005 and re-estimating the empirical model using firm fixed effects.

There is also a concern that the gender-skill division of the firms' labor force may be correlated with unobserved firm- or worker characteristics. Besides the estimation of fixed effects, robustness is assessed by controlling for unobserved and time varying productivity shocks (Olley and Pakes, 1996; Levinsohn and Pethrin, 2003). The influence of unobserved worker effects when testing for equal relative productivities and earnings should, however, be limited if the bias that they introduce in the estimated differentials in earnings and productivities run in the same direction[21].

Estimation of Equations (3) and (4) is conducted simultaneously using the Non-Linear Seemingly Unrelated Regressions (NLSUR) method, which takes account of cross-equation correlation in the shocks to wages and output.

4. Data, variables and summary statistics

This study uses survey data on industrial firms collected by China's National Bureau of Statistics (NBS) in 2004[22]. The dataset covers all state-owned firms and all non-state firms with annual sales above 5 million RMB (about 750,000 US dollar) and includes all firms formally designated as large or medium sized, as well as the bulk of those designated as "small-scale". Dividing the total value added in the dataset by the industrial GDP reported in China's Statistical Yearbook indicates that the dataset accounts for more than 90% of total industrial output.

To ensure high data quality, I follow the cleaning procedure of Jefferson et al. (2008) and omit firms that report zero employment and those with few-

[20] So that $P_{FU} + P_{FS} = 1$ and $P_{MU} + P_{MS} = 1$.

[21] For example, if the unobserved characteristic of tenure is likely to result in an upward bias in both productivity and wages. These biases are then cancelled out when we compare one to the other.

[22] This dataset is used by the NBS to compile the "Industry" section of the China Statistical Yearbook and industry specific reports in the China Markets Yearbook.

er than eight workers, firms in the upper and lower tails of productivity[23], and firms with improbable ratios of value added to productivity[24]. This procedure removes 11,807 firms, leaving a final sample size of 257,721. Details of the variable constructions are presented in section A1 in the Appendix.

Summary statistics are reported in Table 1. They show that the average firm has 248 employees of whom 40% are female. The average share of skilled male employees (8%) is twice the size of the share of female skilled employees (4%). Turning to the ownership-based sub-samples, we see that the share of female employees is lowest in the state-owned sector (33%), somewhat higher in firms belonging to the collective (35%) or private (40%) sectors, and highest in firms with investment from HKMT-based (48%) or foreign entities (49%).[25] State-owned firms are, on average, larger (806 employees) than those in other sectors. The smallest average workforce size is in the collective and private sectors, at around 150 workers. Private domestic firms and Collective owned firms pay the lowest annual wage per worker, and state-owned and Foreign-financed firms pay the highest.

The dataset comprises 44.1% domestic private firms, 8.3% state-owned, 9.6% Collective owned and about 10.5% funded by entities based in Hong Kong, Macao or Taiwan or in foreign countries. The remaining 17% are classified in the "Mixed" ownership category.[26] The bottom row of Table 1 shows the proportion of firms reporting non-zero expenditures on all three non-salary items. Using this definition of social insurance participation, 12% of all firms are found to participate.

The summary statistics for industrial sub-sectors are given in Appendix, Table A3. They are calculated after first grouping firms according to the 14 industry ex-factory price index categorization used in China's statistical publications. The percentage of female workers ranges from 9% in the coal industry to 73% in the tailoring industry.

Figure 1 illustrates the distributions of the proportion of female workers by region, with regions ranked by their level of GDP per capita in 2004. This graphical evidence suggests a positive correlation between the demand for female industrial workers and economic development. Probably, this correlation is closely connected with the geographical variation in the intensity of

[23] After computing ratios of value added to labor and capital, and ratios of labor and capital to value added, we delete firms for which the computed values for each of these four variables is more than four standard deviations above the mean of each measure.

[24] Below zero or above one.

[25] Previous studies suggest that the workforces of export-oriented and Foreign-financed enterprises may be up to eighty percent female (Tan, 2000). In the dataset used in this paper this is true for 11.7% of the HKMT-funded firms, and for 14.6% of the foreign-funded firms.

[26] Jefferson and Su (2006) argue that that formal ownership registration is an unreliable measure of actual firm control. Examining the share of firm assets held by the Chinese state and by non-mainland entities in the different ownership categories however shows that an alternative ownership categorization based on the this variation would only differ slightly from the one based on registration.

market reform, both via the increase in investment from outside the Chinese mainland and via the expansion of light industrial sectors.

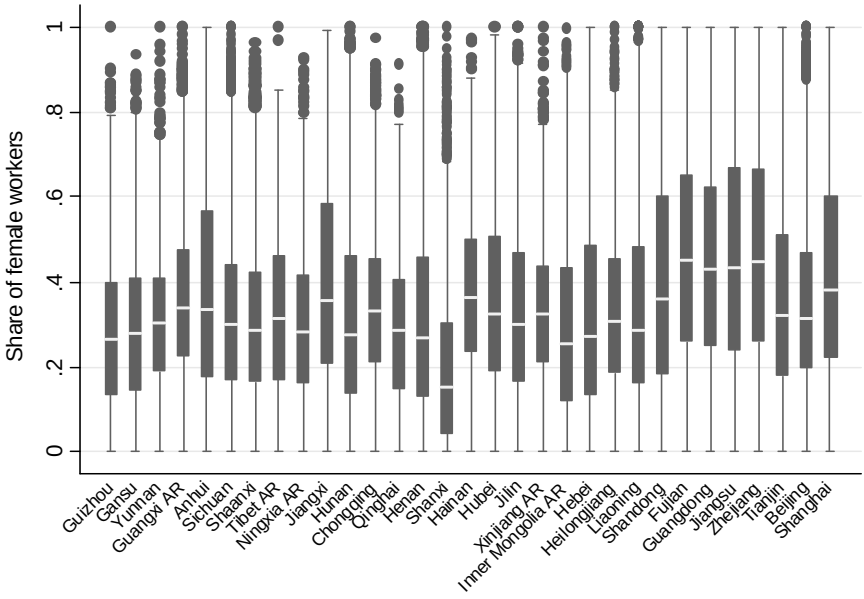

Figure 1: Distribution of the proportion of female employees by region, ranked by 2004 regional GDP per capita

Table 1: Summary statistics.

	All firms	*State-owned*	*Collective owned*	*Private domestic*	*HKMT-financed*	*Foreign-financed*	*Mixed ownership*
N	257,721	21,477	24,732	113,498	26,963	27,139	43,912
(share of total sample)	(100)	(8.3)	(9.6)	(44,1)	(10.5)	(10.5)	(17.0)
Log value added (¥ 1000)	8.75	8.67	8.62	8.42	9.13	9.38	9.10
Log capital (¥ 1000)	8.16	8.95	7.81	7.65	8.54	8.77	8.65
Wage per worker (¥ 1000)	7.07	7.44	6.85	6.65	7.65	7.70	7.32
Employment	248	601	171	129	322	310	344
Proportion of female employees	0.40	0.33	0.35	0.40	0.49	0.48	0.36
Proportion of male unskilled employees	0.52	0.53	0.59	0.53	0.43	0.40	0.53
Proportion of male skilled employees	0.08	0.13	0.06	0.07	0.08	0.12	0.11
Proportion of female unskilled employees	0.36	0.27	0.33	0.37	0.45	0.42	0.31
Proportion of female skilled employees	0.04	0.06	0.02	0.03	0.04	0.06	0.05
Firm size (proportion)							
8-50	0.26	0.25	0.28	0.33	0.14	0.18	0.19
51-100	0.25	0.17	0.27	0.29	0.20	0.21	0.23
101-500	0.40	0.38	0.40	0.35	0.51	0.47	0.44
501-1000	0.05	0.10	0.04	0.02	0.10	0.08	0.08
1001-	0.03	0.10	0.02	0.01	0.05	0.05	0.06
Firm age (proportion)							
< 3 years	0.24	0.69	0.56	0.12	0.26	0.15	0.22
4-7 years	0.24	0.15	0.25	0.22	0.28	0.31	0.24
8-12 years	0.26	0.09	0.12	0.32	0.23	0.27	0.28
> 12 years	0.26	0.07	0.08	0.34	0.23	0.27	0.27
Township or village enterprise	0.17	0.38	0.66	0.00	0.08	0.08	0.35
Proportion of unionized firms	0.46	0.83	0.53	0.36	0.41	0.43	0.58
Social insurance participation (proportion)	0.12	0.39	0.09	0.03	0.12	0.26	0.18
Proportion of social insurance in total wages	0.11	0.31	0.11	0.05	0.09	0.13	0.13

Notes: Social insurance participation is defined as the firm reporting non-zero yearly expenditures for i) unemployment or labor insurance, ii) pension and medical insurance, and iii) housing funds.

5. Are women discriminated against in Chinese industry?

5.1 Baseline estimation results

The results from the nationwide sample of firms are presented in Table 3. Columns 1 and 2 contain estimates derived by jointly estimating the production function (3) and the wage equation (4) for the full sample, with the total wage bill being the dependent variable in the wage equation. Row one contains the productivity and wage differentials between women and men with less than 12 years of education. These estimates show that unskilled women earn, on average, wages that are 20% lower than those of unskilled men, and that they are, on average, 61% less productive. Comparing these two parameters to each other, we find that although women, on average, receive lower wages than men, their productivity lags even further behind. The conclusion, therefore, appears to be that the average Chinese industrial firm prefers to hire unskilled female workers compared to unskilled male workers, a finding that will be more thoroughly discussed below.

Next, we assess how education affects the wages of male and female workers. The average wage estimates in rows 2 and 3 show that workers of both genders are rewarded for their educational attainments, but that women are rewarded more[27]. For both genders, these parameters capture the average wage premium received by a worker with the average education attainment in the "skilled" category compared to a male worker with the average education attainment in the "unskilled" category. To obtain an approximation for the education premium associated with an additional year of schooling, we therefore need to calculate this number of years. Doing so yields an approximate return to return from an additional school year of 12%-15% for men, and 36-46% for women.[28] As we shall see below, this result is because of the

[27] This result is not a consequence of women having a larger average number of years of schooling than men within the skilled worker category. Nor do the results appear biased by the greater on-the-job training given to skilled men compared to skilled women. Although a simple OLS regression shows that firms with one percent more women in their workforces spend on average six percent less on worker education, the benchmark results are robust to estimation in a sub-sample including only the 144,666 firms where those expenditures are zero.

[28] The composition of educational attainment levels among skilled men and women are as follows. Among skilled men: graduate studies (18 years, 2%); undergraduate university (16 years, 25%); professional school "*Dazhuan*" (15 years, 73%). Among unskilled men: high school (12 years, 44%); junior high school or below (5 or 8 years, 57%). Among skilled women: graduate studies (18 years, 1%); undergraduate university (16 years, 18%); professional school "*Dazhuan*" (15 years, 81%). Among unskilled women: high school (12 years, 41%); junior high school or below (5 or 8 years, 59%). The higher and lower "bound" for the calculation is determined by the average years of schooling assumed for workers having attained "junior high school or below", set to lie in the interval 5 to 8 years.

higher wages paid to skilled women compared to skilled men in Collective and State owned firms.

Row 4 contains the calculated gender gaps in average wage and productivity among skilled workers, showing that the higher education reward for skilled women is reflected in an 18% gender wage gap in favor of women. This result mirrors previous evidence on the gender difference associated with the education premium, derived from studies using individual-level data (Maurer-Fazio, 1999; Li, 2003).

Table 2: Joint production function and earnings equation estimates: Cobb-Douglas production function, all firms.

	Production function and wage equation estimates			*Production function and total earnings equation estimates*		
	Log (Wages) (1)	Log (Value Added) (2)	Col. 1 = Col. 2 (*p*-val)	Log (Total Comp.) (3)	Log (Value Added) (4)	Col. 3 = Col. 4 (*p*-val)
Unskilled female	0.80 (.00)	0.39 (.01)	0.42 (.000)	0.79 (.01)	0.38 (.01)	0.41 (.000)
Skilled female	2.22 (.03)	3.64 (.09)	-1.41 (.000)	2.53 (.03)	3.64 (.09)	-1.12 (.000)
Skilled male	1.88 (.02)	3.50 (.06)	-1.61 (.000)	1.96 (.02)	3.48 (.06)	-1.53 (.000)
Skilled female-Skilled male (i)	1.18 (.02)	1.04 (.03)	0.14 (.03)	1.29 (.02)	1.05 (.03)	0.25 (.03)
Female-male (ii)	0.88 (.01)	0.57 (.01)	0.28 (.000)	0.89 (.01)	0.57 (.01)	0.32 (.000)
Log labor		0.64 (.00)			0.64 (.00)	
Log capital		0.20 (.00)			0.19 (.00)	
Rural admin. subordination	-0.13 (.00)	-0.11 (.01)		-0.16 (.00)	-0.11 (.01)	
R-squared	0.84	0.60		0.84	0.60	
R-squared between equations	0.12			0.13		
N	255,568			255,733		

Notes: All equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), industry (38), size (5), age (4), ownership (6), unionization, and township-and-village enterprise status.

(i) Wage and productivity differentials for skilled women vs. skilled men, and

(ii) skill-weighted wage and productivity differentials for all women vs. all men.

Comparing the magnitude of the private return from education and the firm's return from hiring an educated worker, I find that the payoffs for education received by both women and men fall short of the value of their contributions to the firm's output. While this result mirrors those of previous studies (e.g. Hellerstein et al., 1999; Hellerstein and Neumark, 1999, 2005; Dong and Zhang, 2009), it remains puzzling. Fleisher et al. (2004; in press) argue that a likely explanation in the Chinese case could be inefficiencies in competition and in the allocation of labor and capital. In terms of competition, State-controlled firms in particular may have substantial monopsony power, which allows them to under-compensate skilled workers in labor markets that fail to achieve a competitive equilibrium. For non-state firms, constraints on borrowing, land-use and electricity (power) may preclude them from achieving their profit-maximizing size[29]. Moreover, obstacles to labor mobility stemming from the family registration (*hukou*) system could prevent the emergence of a competitive market for skilled labor.

Finally, row 5 of Table 3 presents the average skill-weighted wage- and productivity differences between all women, unskilled and skilled, compared to all men. From an international perspective, the narrow wage-gap is particularly noteworthy. Using similar data and methodologies, Hellerstein and Neumark (2005) report that women earn 38% less than men in the US; Bartolucci and Alberto (2010) find a 34% gender wage gap in the case of Germany; and Asano and Kawaguchi (2007) find a 70% gap in the case of Japan. In comparison, the 22% gap found in this study is quite small and may suggest that, at least on average, the Maoist ideology of equal pay has been maintained during the economic transition.

5.2 Does non-wage compensation contribute to a wider gender-gap in earnings?

To examine whether non-wage compensation provided by firms contributes to a wider gap in gender earnings, the sum of the firms' expenditures on social insurance payments is added to the wage measure in Equation (4) before joint estimation with the production function. Using the entire sample of firms, out of which a large proportion do not participate in the social insurance programs, I first examine the combined effect of women receiving fewer benefits within firms, and the potential over-representation of women within non-participating enterprises. The results from this exercise are shown in the right-hand half of Table 3. In examining the results, I am interested in comparing the gender gap measured in terms of total earnings (Column 3) to the benchmark wage gap presented in Column 1. The comparison

[29] Poncet et al. (2009) use data for 20,000 Chinese manufacturing firms from 1998-2005 to show that private domestic firms are credit constrained.

shows that the gender gap for total earnings is very similar to that of wages (11% compared to 12%) and that the difference is not statistically different from zero.

Examining the earnings gaps for unskilled and skilled workers separately, I cannot detect any gender polarization of non-wage earnings for unskilled workers. In contrast, I find that the earnings advantage of skilled women compared to skilled men grows from 18% to 29% when taking account of non-wage compensation. Further examination of this result is carried out by omitting firms with zero contributions for all three social security items. In this smaller sample I do not find a statistically significant difference between the gender gap with respect to wages or total earnings for the skilled worker group. Hence, rather than receiving more non-wage compensation within firms, the aggregate result appears to stem from a greater tendency for skilled women, compared to skilled men, to be working in firms that comply with the social insurance regulations.

Examining the returns from education, measured in terms of total earnings and presented in rows 2 and 3 of Column 3, we can see that non-wage compensation appears to be distributed slightly in favor of workers with higher levels of education. Compared to the education payoff in terms of wages, the average additional income of skilled women compared to unskilled men increases by 31 percentage points measured in terms of total earnings. That of skilled men increases by 8 percentage points. This observation is in accordance with incentives for firms to reduce turnover costs by providing skilled workers with social insurances which were, at least in 2004, generally not transferable between workplaces.

We next exclude employers that do not participate in the social insurance programs from the sample, leaving 31,554 firms. For these employers, the average ratio of non-wage benefits to wages is 0.37. As shown in Panel A, Table A4, the results for this sample are not different from those for the whole sample. Including social insurance payments in the earnings measure does not yield a wider gender gap in work compensation.

In 2004, China's social insurance system was highly fragmented. In many regions workers with temporary contracts (often migrants) lacked rights to insurance cover, or were entitled to inferior insurance packages than permanent workers. Several robustness checks are carried out to ascertain whether the small effect of social insurances on the baseline gender gap in wages is not caused by an underestimation of the total earnings bill. In Table A4, Panel B, I first use data from Guangdong province only (32,545 observations). Since the late 1990s, this province has taken the lead in requiring firms to provide equal insurance cover for temporary and permanent contract workers (Zhang et al., 2010). In Panel C, I exclude the most likely employers of migrant labor, namely private firms, from the sample of participating firms included in Panel A. Results from both these sub-samples corroborate previous findings. There is no statistically significant difference between the

gender gaps in wages and the gender gaps in total earnings for either unskilled or skilled workers.[30]

5.3 Robustness checks

I first check the robustness of the average gender differences in wages and productivity with respect to the inclusion of firm fixed effects. Relying on variations within firms rather than across firms, I can check the baseline results for biases associated with unobserved firm or worker characteristics and systematic sorting of male and female workers across firms. Adding data for the year 2005, we obtain a sample of 185,508 firms that were operating in both years.[31] The 2005 dataset, however, lacks information on the human capital distribution of the firms' workforces, and I can thus only calculate average wage and productivity differentials that are not weighted for worker skills.

Using the pooled dataset for 2004 and 2005, I expect the removal of the grouping of workers as skilled or unskilled to yield narrower gender gaps in productivity and wages. This is precisely the finding when estimating the unweighted gaps using the 2004 sample for comparison. The results show that women's average wages are 82% of men's average wages ($\lambda_F = 0.82$), and that their productivity is 75% of men's ($\phi_F = 0.75$). When including fixed effects in the two-year panel, the gaps narrow even further, indicating that sorting may be of at least some relevance. Women's wages are now 92% of men's ($\lambda_F = 0.92$), and their productivity 85% of men's ($\phi_F = 0.85$)[32]. Both these gaps are statistically different at the one percent level. Subtracting one from the other indicates an average over-compensation of women compared to men of 7% (*p-value: 0.000*).

A common criticism of production function estimations is that unobserved productivity shocks can bias the input parameters. The estimated coefficients for capital and labor in the benchmark production function estimates in Table 3 are both highly significant and of magnitudes that seem plausible. Nevertheless, they may be biased if the firm's input choices are correlated to unobserved shocks to productivity. In particular, if female labor adjusts more speedily than male labor to unobserved productivity shocks, there could be an upward bias in the relative wage and productivity estimates for women.

[30] The same result is found when firms located in the Eastern region (Beijing, Tianjin, Hebei, Liaoning, Shanghai, Jiangsu, Zhejiang, Fujian, Shandong, Guangdong, and Hainan) are excluded, as well as when firms with an above-median share of unskilled workers are excluded.

[31] Monetary variables are deflated to 2004 prices using Ex-factory price indices reported in China's Statistical Yearbook.

[32] The coefficients for capital and labor in the production function are 0.31 for and 0.69 respectively. Both are statistically significant at the one percent level.

In addition to the fixed effect approach taken above, we control for unobserved shocks to productivity by using the method proposed by Olley and Pakes (1996) and extended by Levinsohn and Petrin (2003). These methods hinge on the assumption that labor is a less flexible input than capital, and that we can use a polynomial of capital and investment (OP), or capital and materials (LP)[33], to account for the shocks.[34] The results presented in the first two columns in Table A5 support the baseline findings of the paper, since the estimates of relative female wages and productivity are almost identical to the results without proxy variables.

The lack of information about hours worked could be a potential source of error in the quality of labor index. A headcount variable of the female and male workers per firm would fail to reflect gender differences in the supply of labor if the average number of hours worked varies systematically by gender. Using individual-level data from the 2002 China Household Income Project (CHIP), labor input weights are calculated for each gender-skill group and for each ownership sector[35]. The usual number of hours worked in a week is multiplied by the number of weeks worked in a month and the number of months worked in a year. The weights are created by normalizing the average yearly working times for each gender-skill group with the calculated working times for unskilled male workers. Examining the weights in Table A6 confirms some basic expectations about gender and skill patterns in relation to work time. Employees with higher levels of education work fewer hours per week. Moreover, female unskilled employees are found to supply less labor than male unskilled employees in SOEs, while the reverse is observed in Foreign-financed firms. Panel A in Table A7 presents the results of estimations derived both with and without labor input weights. They show that the general conclusions associated with the previous results are not altered.

It has been argued that technological change is gender-biased due to improvements in firm productivity leading to the dismissal of (generally male) production labor (Berman et al., 1994; Dunne et al., 1997). Under such circumstances, the productivity of females in our analysis would be overestimated due to the positive correlation between the proportion of women and productivity enhancing technological change. Panel B in Table A7 examines this hypothesis by dividing the data into two sub-samples based on the median proportion of women. If the hypothesis is true, controlling for the pro-

[33] Besides increasing data availability, Levinsohn and Petrin (2003) also argue that replacing the investment variable with materials is appropriate if investment is lumpy or carries adjustment costs. For both the OP and the LP methods, we follow Olley and Pakes in expanding the Cobb-Douglas production function with the third degree polynomial.

[34] Given the firm's capital stock, its investment is assumed to be a monotonic function of the firm-specific productivity shocks. This investment function is then inverted to yield the unobserved state variable as a function of capital and investment decisions.

[35] The HKMT and Mixed ownership sectors are however excluded due to a lack of unambiguous ownership information for these sectors in the CHIP data.

portion of females by splitting the sample in this way should yield larger gender differences in productivity. Comparing the estimation results within the two sub-samples to the baseline in Table 2, however, shows that the skill-weighted average gender-difference in productivity remains roughly constant.

If women take jobs in more labor intensive firms and labor contributes a decreasing return in the firm's production function, the proportion of women would be negatively correlated to firms' productivity. I test this hypothesis by dividing the entire sample on the basis of the median capital-to-labor ratio. The estimation results do not provide support for the idea that women are sorted into more labor intensive firms with low labor productivity. The productivity differential between men and women (Panel C, Table A7) remains largely unchanged despite the sample division.

In China, labor unions are responsible, by law, for monitoring firms' compliance with anti-discriminatory laws and regulations. If unionization is correlated with the proportion of male workers, men's wages could receive a higher degree of protection than women's. Empirical research has provided evidence of higher wages, non-wage benefits and productivity in Chinese firms with labor unions (Ge, 2007). This potential bias is assessed by comparing estimation results between the sub-sample of 119,788 firms with labor unions, and the 137,933 firms without unions (results not reported). No statistically significant differences with the benchmark findings are found for either group.

6. Extensions

6.1 Firm ownership

Table 3 presents results obtained by jointly estimating Equations (3) and (4) for six ownership sectors: 1) State-owned, 2) Collective owned, 3) Privately owned, 4) firms funded by entities based in Hong Kong, Macao or Taiwan, 5) Foreign-financed firms, and 6) a residual category of firms with so-called "Mixed" ownership forms.

The ownership-specific results shed light on some interesting dynamics in relation to the gender gaps in wages, as well as in employers' preferences for skilled and unskilled female labor. Consider first our group of skilled workers. The benchmark result that women receive larger average education premiums is not constant across ownership types. Highly educated women earn higher wages than highly educated men in State-owned firms, Collective owned firms and Private domestic firms. Meanwhile, there is a small and statistically non-significant gender wage gap among skilled workers in firms with investments from foreign or HKMT-based sources. Consider instead

the pattern of reward between men and women with limited education, who make up 87% of the workforce in the average firm. For these workers, the gender pattern in reward shows more consistency across sectors: there is a statistically significant wage disadvantage for women with limited education regardless of firm ownership. The narrowest wage gap is in the Private domestic sector (13%) and the widest in State owned firms (28%). Thus, the results appear to contradict the notion of expanding gender wage gaps in the industries established during the market reform period in comparison to older State-owned firms.

The results conform to those presented in previous literature and that used the CHIP household data to compare gender wage gaps between skilled and unskilled workers (Millimet and Wang, 2006; Hughes and Maurer-Fazio, 2002; Xu et al., 2006; Gustafsson and Li, 2000; Bishop et al., 2005). Because State-owned firms are over-represented in the CHIP data and Foreign-financed ones under-represented, the previous results uning this dataset is comparable to my subsample of the State-owned sector. In this sample, the results mirror those of previous studies in identifying a narrower gender-wage gap among unskilled than among skilled employees.

The calculated skill-weighted average gender wage gaps by ownership sector also corroborate previous research findings. Studies focusing on ownership, rather than education level, have shown larger-than-average wage differences between men and women employed in the Foreign-financed sector (Maurer-Fazio et al. 1999; Hughes and Maurer-Fazio, 2002; Liu et al., 2000; Xu et al., 2006). This finding is echoed in my analysis as I find an aggregate gender wage gap of 16% in the Foreign-financed firms and 21% in firms with investment from Hong Kong, Macao or Taiwan, while the gap is somewhat smaller in State-owned firms (12%), Private firms (10%) and Collective owned firms (8%). Furthermore, the results indicate that the pattern of wider gaps in sectors with investment from outside mainland China stems from two sources: i) wider gender wage gaps among unskilled workers, and ii) the lack of a wage-advantage for highly educated women compared to highly educated men.

I now compare how the relative wages of well- and poorly-educated women match their average productivity across the company ownership types. For women with limited education, the results demonstrate that wages exceed productivity in all sectors. For highly educated women, average wages match average productivity in all sectors except State-owned firms and Mixed ownership firms. In the case of the State-owned firms, the result is consistent with the reasoning of Dong and Zhang (2009). They argue that the legacy of the planned economy, which involved disregarding the average lower physical capacity of female workers while maintaining gender-equal wages, resulted in a systematic overcompensation of women in comparison to their productivity contributions.

Table 3: Joint production function and wage equation estimates: Cobb-Douglas production function, by ownership.

	State-owned			*Collective owned*		
	Log (Wages) (1)	Log (Value Added) (2)	Column 1 = Column 2 (*p*-value)	Log (Wages.) (3)	Log (Value Added) (4)	Column 3 = Column 4 (*p*-value)
Unskilled female	0.72 (.03)	0.48 (.05)	0.24 (.000)	0.86 (.01)	0.30 (.03)	0.56 (.000)
Skilled female	2.44 (.11)	3.46 (.28)	-1.02 (.000)	1.99 (.10)	3.48 (.38)	-1.51 (.000)
Skilled male	2.18 (.07)	3.95 (.21)	-1.81 (.000)	1.42 (.05)	2.71 (.21)	-1.32 (.000)
Skilled female-Skilled male (i)	1.12 (.05)	0.87 (.08)	0.25 (.002)	1.40 (.10)	1.28 (.19)	0.12 (.530)
Female-male (ii)	0.88 (.03)	0.70 (.05)	0.18 (.000)	0.92 (.02)	0.49 (.04)	0.43 (.000)
N	20,830			24,532		
	Foreign-financed			*HKMT-financed*		
Unskilled female	0.78 (.02)	0.42 (.02)	0.36 (.000)	0.75 (.01)	0.43 (.02)	0.32 (.000)
Skilled female	2.65 (.09)	4.47 (.27)	-1.82 (.000)	2.10 (.08)	3.57 (0.24)	-1.47 (.000)
Skilled male	2.77 (.07)	5.00 (.25)	-2.21 (.000)	2.15 (.06)	3.96 (.21)	-1.81 (.000)
Skilled female-Skilled male (i)	0.95 (.04)	0.90 (.04)	0.05 (.348)	0.98 (.05)	0.90 (.05)	0.07 (.336)
Female-male (ii)	0.84 (.02)	0.67 (.03)	0.17 (.000)	0.79 (.01)	0.57 (.03)	0.21 (.000)
N	27,024			26,893		
	Private Owned			*Mixed Ownership*		
Unskilled female	0.87 (.01)	0.39 (.01)	0.48 (.000)	0.79 (.01)	0.35 (.02)	0.44 (.000)
Skilled female	1.70 (.04)	3.50 (.17)	-1.80 (.000)	2.30 (.07)	3.54 (.23)	-1.23 (.000)
Skilled male	1.45 (.02)	2.89 (.09)	-1.44 (.000)	1.89 (.04)	3.43 (.14)	-1.51 (.000)
Skilled female-Skilled male (i)	1.17 (.04)	1.21 (.08)	-0.04 (.616)	1.21 (.05)	1.03 (.08)	0.18 (.029)
Female-male (ii)	0.90 (.01)	0.57 (.02)	0.37 (.000)	0.90 (.02)	0.61 (.03)	0.28 (.03)
N	112,859			43,547		

Notes: All equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), industry (38), size (5), age (4), unionization, R&D expenditure per worker, township-and-village enterprise status.

(i) Wage and productivity differentials for skilled women vs. skilled men, and

(ii) skill-weighted wage and productivity differentials for all women vs. all men.

How are we to understand the finding that firms in the Private and Foreign-financed sectors also pay unskilled women wages that seem to exceed their contributions to firm productivity? Examining the qualitative research on gender and production in China's newly established industrial sectors, it appears likely that this situation stems from a preference for female labor within a gender-segregated manufacturing organization. The "feminization of labor" is a standard observation in rapidly internationalizing and emerging economies (Standing, 1989; 1999; and e.g. for Mexico by Salzinger, 2003). In the Chinese case, this type of feminization in production is mirrored by the reliance on unskilled women in assembly line work, in particular in export-processing firms that sell light manufacturing goods on the international market.

Qualitative studies of Chinese factories have shown that workforce control constitutes a major reason for employers' preference for unskilled female labor over unskilled male labor. In the eyes of employers, women are perceived as more docile, obedient and easily controlled (Lee, 1995; Pun, 2005; Xue, 2008). Control is further enforced via the use of single-sex dormitories, which, together with gender segregated roles in the production process, Pun and Smith (2006) call the "dormitory labor regime". Besides the desire to reduce costs via increased workforce control, it is likely that, once the separation of unskilled women and men on the shop floor and in their living accommodation is established, there would be increased costs if the model is departed from. As such, the motives of workforce control, combined with gendered organization of production and living spaces within firms, can explain the observed preference for unskilled female labor in privately controlled industrial enterprises.

6.2 Industrial segregation

If women are systematically hired by industries that pay less and have lower productivity, such sex-segregation may cause an upward bias in the estimates of women's relative wages and productivity over the whole sample. Returning briefly to the descriptive statistics subdivided by industry, presented in Table A3, the female labor force indeed appears to be systematically divided across sectors. Women make up less than a third of the workforce in the metallurgical, petroleum, power, timber and coal industries, while they make up more than two thirds in the tailoring and textile sectors.

The results from estimating Equations (3) and (4) for each of the 14 aggregated industrial sectors are presented in Table A8. An examination of the results gives little support for a strong effect of sex-segregation on the estimates. For unskilled workers, a wage disadvantage for women greater than the benchmark result of 20% is found in eight of the fourteen sectors, and a

gap slightly narrower than 20% is found in three of them. Only in the Tailoring, Paper, and Chemical industries, was no statistically significant wage-difference found. For skilled workers, there is substantial variation in the size of the gender wage gap across the sectors[36]. Women's payoffs from education are found to be larger than men's in seven sectors, smaller in the Power and Petroleum sectors, and on a par with men's in five highly labor intensive industries[37].

There is a clear productivity disadvantage associated with unskilled women across all 14 industrial sectors. In addition, unskilled women consistently have average wages that surpass their average productivity. It is noteworthy that the latter result does not only apply to heavy industrial sectors, but also to light manufacturing. As such, it yields further support for the hypothesis that there is an employer preference for unskilled women in light and labor intensive industries. As argued above, that preference is probably rooted in the production organization of assembly style manufacturing plants, in which tasks tend to be divided by gender, with unskilled women working at the conveyer belts on the shop floor. In addition, there could be fixed and unobservable costs associated with hiring unskilled men instead of unskilled women. For example this would be the case when the firm provides separate living quarters for the two genders.

For the heavy industrial sectors, the "preference" for female workers, both unskilled, and in some sectors also skilled[38], adds support to Dong and Zhang's (2009) conjecture regarding the legacy of a mismatch between female workers and jobs.

Finally, it is noteworthy that the smallest estimates of the gender gaps in wages are in the more labor intensive and recently established industrial sectors. These sectors are also the ones that have the largest average share of female employees. Although this result could be interpreted as women faring relatively better in these sectors, it is worth remembering that wages are only one out of many aspects associated with a job. As argued in previous studies, women in the "new" industrial economy may be systematically hired to fill the more precarious jobs with poor job security (Razavi, 2007; Cooke, 2001), and may face exploitative working conditions (Burda, 2007; Pun, 2007; Berik et al., 2007).

[36] The results for the Coal industry are however probably spurious. Firms in this industry are outliers with respect to the dominance of State ownership, firm age, and the low average proportion of female workers (9%) and skilled workers (4%).

[37] These results thus highlight the importance of industry representativeness of the data sample when assessing gender differences in the returns from education among Chinese industrial workers.

[38] Over-compensation of skilled women compared to their average productivity is observed in the Timber, Power, and Food sectors, all of which are sectors with firms that are older than average and more likely to be State-controlled.

6.3 Regional variation in market reform intensity

China's strategy of market reform has followed a distinct geographical pattern. Starting with a few select Special Economic Zones on the East coast, the liberalization of labor and product markets was gradually extended first to wider coastal areas, and then to inland and western locations. Numerous studies provide empirical support for a strong effect of the spatially diverse liberalization policy on economic growth (Chen and Feng, 2000), even when accounting for geography (Démurger et al., 2002), domestic investment (Jones et al., 2003) and agglomeration (Chen, 2009).

In an attempt to assess the effect of reform intensity on women's relative wages and productivities, I estimate the wage and production functions for each of China's regions (22 provinces, 4 self-governing municipalities, and 5 Autonomous Regions). Regions are ranked by their level of GDP per capita in 2005 to examine the indirect effect of market reform. Four regions with less than 1,000 observations are excluded from the analysis[39].

To limit space I only report the results for workers with less than twelve years of education. Figure 2 shows the estimated gaps in wages and productivity by region, with the regions ranked according to their GDP per capita in 2004. The horizontal fitted lines indicate that there is no connection between reform and women's relative wages or productivity[40]. This result mirrors the finding of Xie and Hannum (1996) who combine 1995 CHIP data with city-level variations in industrial growth as a proxy for market reform and conclude that the size of the gender earnings-gap is unrelated to industrial output.

In Figure 2, the gap between the trend lines of the estimated wage and productivity gaps can arguably be understood to represent the evolution of wage discrimination experienced by unskilled women during the economic transition. We see that, on average, women are paid in excess of their contribution to firms' productivity in all regions, and that the gap is similar across regions. This observation appears to run counter to the expectation that competition decreases the possibility of employers indulging in discriminatory wage-setting (Becker, 1971). However, and as noted above, economic reform may have triggered the growth of industrial sectors in which firms gained a competitive edge by organizing production on the basis of gender segregation. Thus, a preference for docile female assembly line workers housed in single-sex dormitories could have extended the "over-payment" of unskilled women despite market reform.

[39] Qinghai (391), Ningxia Automonous Region (597), Tibet Autonomous Region (158), and the island of Hainan (513)

[40] Both correlations are also statistically non-significant.

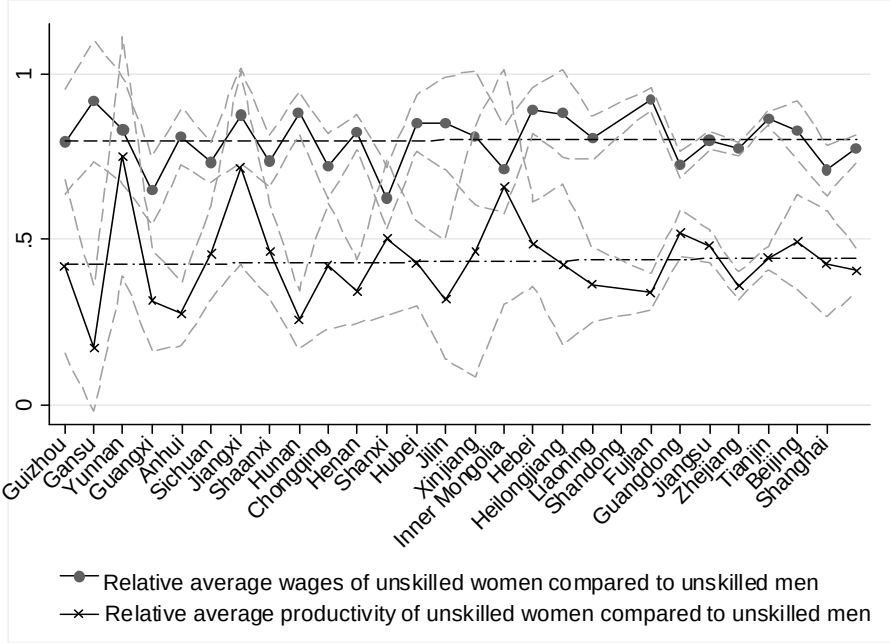

Figure 2: Average wages and productivities of unskilled women relative to unskilled men by region, ranked according to regional GDP per capita.

7. Conclusions

This paper used firm-level data from Chinese industrial firms to assess differences between men and women's work compensation relating to education level, ownership of the firms, the firm's industrial sector, and its location. The paper also investigated whether non-wage payments in the form of social insurance premiums contributed to a wider gender gap in earnings compared to wages when considered in isolation.

Estimation based on a highly representative sample of Chinese firms showed that women earned on average 18% less than men. This gender wage gap is narrower than those found for other large economies in studies using firm-level data and similar methodologies. Thus, the result could suggest that the Maoist ideology of gender equality carried over into the period of market transition. Contrary to the view of the State-sector being a protector of women's rights, the average gender wage gap was not narrower in state-owned firms than in other ownership sectors.

Investigating the role of social insurance payments in the gender gap in earnings showed that these payments were not disproportionate with respect to the male workers. For workers with more than twelve years of education, the results actually demonstrated a greater tendency for women to be employed in firms abiding by the regulations, translating into a slightly greater

provision of social insurances for skilled women compared with skilled men. These results indicate that the fears of gender discrimination with respect to provisions of pensions, medical care, housing funds and unemployment insurances by employers could be exaggerated.

In accordance with previous studies of the importance of education level for the gender-wage gap in Chinese industry, I find that women with a lower level of education have a wage disadvantage compared to men with limited education, but that women with a higher level of education earn higher wages than highly educated men. In this respect, the ownership and industry specific estimations demonstrated that the wage advantage of skilled women was driven by the wage-setting practices in State-controlled sectors and heavy industry.

For women with less than twelve years of education, their productivity disadvantage was found to be larger than their wage disadvantage, a result that was consistent across industrial, regional, and ownership groups. For heavy industrial firms, this result could be ascribed to a legacy of mismatching women with jobs during the planned economy, as argued by Dong and Zhang (2009). For more recently established light industrial and foreign-owned sectors, the persistence of a positive wage-premium for unskilled women is contrary to Becker's (1971) prediction that competitive forces should push firms to improve efficiency by paying workers in accordance with their productivity contributions. Instead, the results appear more in line with the gender segregated production processes observed by qualitative researchers (e.g. Pun, 2005). This research has identified an employer preference for female workers, presumably motivated by their perceived obedience and docility. In addition, a preference for females can stem from observed gender diversification of firms' production processes. In such settings, additional fixed costs for hiring a man instead of a woman can arise if firms follow the policy of hiring only women for assembly line jobs, and if men and women are housed in separate company-supplied dormitories.

The observed preference for female workers in more recently established industries should however be considered in the light of the continuing productivity disadvantage of unskilled women in these sectors. The identification of a productivity disadvantage associated with female industrial workers is a standard result in empirical studies using firm-level data, including when using employer-employee datasets and incorporating advanced controls for sorting; the reasons for this gender difference, however, remain elusive. More research is needed to develop an understanding of these mechanisms in order to take advantage fully of the productivity potential of the whole industrial workforce.

Finally, the results provide some interesting contrasts to studies of labor market developments in other transition economies. In these countries, the gender wage gap either remained constant or became wider during the transition, and the demand for female labor declined (e.g. Brainerd, 1998). In the

Chinese case, the results of this paper showed both a more narrow gender wage gap, and an increase in the demand for female labor, in newly established light industrial sectors. Allowing light and export-oriented industries to grow in this way appears to have had positive effects on the Chinese women's labor market.

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Appendix A

Table A1: Components of China's urban wage (gongzi) or earnings statistics for on-post "staff and workers".

Monthly or annual salary income (including base earnings and additions based on position, seniority, wage scale, etc.)
Earnings during on-the-job training, probationary period
Employee income paid on an irregular basis
Hourly payment for work performed
Piecework payment for work performed
Bonus payments
Incentive, performance-based payments
Overtime pay
Hardship, danger pay
All kinds of subsidies in cash or in kind
Festival, holiday subsidy
Travel money, food allowance while traveling
Personal services such as baths, haircuts
Books, newspapers, magazines provided for employees
Meals provided, food allowance
Transport subsidy (car or shuttle bus provided, cash for bus or taxi, and so on)
Housing subsidy (dormitory provided, or directly subsidized rent or purchase of housing)
Individual income tax deducted from earnings and paid directly by enterprise to government
Social insurance funds (pension, medical, unemployment insurance funds, and housing purchase fund) deducted from the employee's wage and paid by the work unit to the government on behalf of the employee
Money for rent, and utilities (electricity, water)
Money for fixed line or mobile phone
Clothing subsidy
Subsidy compensating workers for lack of vacation time
Earnings during approved leaves of absence, pay for time not worked (regular vacation, compassionate leave, to visit relatives, family-planning operation, national or societal duty, study leave, leave due to sickness or injury)
Anything that has the nature or spirit of labor earnings, even if it is not spelled out in the regulations

Source: *Laodong gongzi; tongji taizhang* [*Labor wages; statistical accounts*] (Beijing, Beijing Municipality Statistical Bureau, 2004), pp. 2–1 to 2–5, cited in Banister (2005)

Table A2: Aggregation of the registration-based classification of firms into ownership categories.

Registration-based classification (since 1998)	Categories under which data on all industrial enterprises are reported in the Statistical Yearbook
Domestic enterprises	
State-owned	State
Collective owned	Collective
Employee shareholding company	Collective
Joint operation enterprises	
State-owned	State (1)
Collective owned	Mixed
State- and Collective owned	Mixed (2)
Other joint operation enterprises	Mixed (2)
Limited liability companies	
Solely state-owned	State
Other	State or Mixed (3)
Stock companies	State or Mixed (3)
Private enterprises	
Private sole proprietorships	Private
Private partnerships	Private
Private limited liability company	Private
Private stock companies	Private
Other enterprises	Mixed
HKMT-financed enterprises	
Joint equity ventures (JVEs)	HKMT or State (2)
Contractual joint ventures (CJVs.)	HKMT or State (2)
Wholly HKMT-owned	HKMT
HKMT stock companies	HKMT or State (2)
Foreign-financed enterprises	
Chinese-foreign JEVs.	Foreign or State (2)
Chinese-foreign CJVs.	Foreign or State (2)
Wholly foreign-owned	Foreign
Foreign-financed stock company	Foreign or State (2)

Notes: The table details the ownership aggregation of the registration-based firm classification system into six broader categories when reporting data on all industrial firms in the China Statistical Yearbook. Departures from the CSY aggregation method are reported in footnotes (1)-(3) below. Information about the CSY aggregation methodology is contained in Holz and Lin (2001).

(1) Unlike in the Statistical Yearbook, this category is not double-counted as "Mixed" (here termed "Mixed").

(2) When ownership-disaggregated statistics are reported per ownership in the Statistical Yearbook, a proportion of the statistic is double-counted in the "State" category. This proportion corresponds to the aggregate share of the sum of state capital to the sum of total paid-in capital minus individual capital in the registration-based classification category. I instead count an individual firm as State-owned if the state's share in total capital is greater than 50%.

(3) This category is counted as "State" in the Statistical Yearbook for firms that are under absolute state control (*guoyou juedui konggu*) or relative state control (*guoyou xiahgdui konggu*). The first implies that the state owns for more than 50% of the total capital. The second that the state holds less than 50% of total capital but that i) its share is relatively large compared to the shares of other ownership categories, or ii) even though one or more other ownership categories have a larger capital share, the state in effect holds the control rights by agreement (*Xiyi kongzhi*). In this paper, only absolute state-controlled firms can be identified, and these are moved from the "Mixed" to the "State" category.

Table A3: Summary statistics by industrial sector.

	Metallurgical	*Power*	*Coal*	*Petroleum*	*Machine Building*	*Building Materials*	*Timber*
N	29,939	7,678	4,963	2,427	37,691	68,873	18,547
Log value added (¥ 1000)	8.79	9.17	9.22	9.56	8.74	8.77	8.80
Log capital (¥ 1000)	8.01	10.21	8.56	9.31	8.22	7.99	8.52
Log wages (¥ 1000)	6.88	7.68	7.68	7.24	6.87	7.19	7.09
Employment	237	387	849	732	186	244	214
Proportion of female employees	0.25	0.32	0.09	0.26	0.38	0.36	0.29
Proportion of skilled	0.09	0.23	0.04	0.19	0.15	0.17	0.09
Age	8.45	20.17	14.65	8.93	9.10	9.95	10.33
Union	0.41	0.87	0.58	0.56	0.46	0.47	0.51
Proportion of foreign and HKMT funded firms	0.13	0.05	0.00	0.10	0.21	0.24	0.12
Proportion of State- and Collective owned firms	0.19	0.71	0.52	0.26	0.16	0.17	0.21

	Chemical	*Food*	*Textile*	*Tailoring*	*Paper*	*Culture and Education*	*Leather*
N	7,495	20,574	22,825	11,506	6,016	11,735	8,052
Log value added (¥ 1000)	8.46	8.79	8.65	8.67	8.72	8.43	8.58
Log capital (¥ 1000)	7.80	8.32	8.15	7.54	7.61	8.23	7.56
Log wages (¥ 1000)	6.86	6.64	7.08	7.54	7.42	6.80	7.28
Employment	188	190	246	282	342	159	276
Proportion of female employees	0.39	0.40	0.66	0.73	0.59	0.42	0.59
Proportion of skilled	0.07	0.14	0.05	0.06	0.05	0.10	0.07
Age	6.73	9.40	7.34	7.22	7.37	11.17	8.07
Union	0.34	0.45	0.43	0.41	0.42	0.46	0.42
Proportion of foreign and HKMT funded firms	0.25	0.18	0.22	0.43	0.40	0.16	0.41
Proportion of State- and Collective owned firms	0.08	0.16	0.08	0.06	0.06	0.23	0.09

Table A4: Joint production function and total compensation equation estimates: Cobb-Douglas production function.

	Production function and wage equation estimates			*Production function and total earnings equation estimates*		
	Log (Wages) (1)	Log (Value Added) (2)	Column 1 = Column 2 (*p*-value)	Log (Total Earnings) (3)	Log (Value Added) (4)	Column 3 = Column 4 (*p*-value)
Panel A: Participating firms						
Unskilled female	0.65 (.02)	0.36 (.03)	0.29 (.000)	0.63 (.01)	0.36 (.03)	0.27 (.000)
Skilled female	2.43 (.07)	3.45 (.19)	-1.03 (.000)	2.48 (.07)	3.43 (.18)	-0.95 (.000)
Skilled male	2.33 (.05)	4.34 (.18)	-2.01 (.000)	2.24 (.05)	4.28 (.18)	-2.05 (.000)
Skilled female-Skilled male (i)	1.04 (.04)	0.80 (.05)	0.24 (.000)	1.11 (.04)	0.80 (.05)	0.30 (.000)
Female-male (ii)	0.84 (.02)	0.64 (.03)	0.20 (.000)	0.85 (.02)	0.64 (.03)	0.21 (.000)
N	31,552			31,567		
Panel B: Firms in Guangdong Province						
Unskilled female	0.80 (.01)	0.48 (.02)	0.32 (.000)	0.79 (.01)	0.48 (.01)	0.31 (.000)
Skilled female	2.71 (.09)	4.05 (.27)	-1.34 (.000)	2.94 (.10)	4.05 (.27)	-1.11 (.000)
Skilled male	2.20 (.05)	3.82 (.19)	-1.62 (.000)	2.25 (.06)	3.80 (.19)	-1.55 (.000)
Skilled female-Skilled male (i)	1.23 (.06)	1.06 (.09)	0.17 (.071)	1.30 (.06)	1.07 (.09)	0.23 (.013)
Female-male (ii)	0.87 (.01)	0.64 (.03)	0.23 (.000)	0.89 (.02)	0.64 (.03)	0.24 (.000)
N	32,545			32,548		
Panel C: Participating firms, excluding private domestic firms						
Unskilled female	0.63 (.02)	0.32 (.02)	0.31 (.000)	0.61 (.02)	0.32 (.03)	0.29 (.000)
Skilled female	2.53 (.08)	3.52 (.23)	-0.99 (.000)	2.63 (.08)	3.50 (.23)	-0.87 (.000)
Skilled male	2.44 (.06)	4.48 (.20)	-2.03 (.000)	2.36 (.05)	4.43 (.20)	-2.07 (.000)
Skilled female-Skilled male ii)	1.03 (.04)	0.79 (.06)	0.25 (.000)	1.12 (.04)	0.79 (.06)	0.32 (.000)
Female-male (ii)	0.83 (.02)	0.62 (.04)	0.21 (.000)	0.86 (.02)	0.63 (.04)	0.23 (.000)
N	28,196			28,209		

Notes: All equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), industry (38), size (5), age (4), ownership (6), unionization, R&D expenditure per worker, township-and-village enterprise status. Geographical location is excluded in Panel C.

(i) Wage and productivity differentials for skilled women vs. skilled men, and

(ii) skill-weighted wage and productivity differentials for all women vs. all men.

Table A5: Additional production function estimations.

	Olley Pakes (1)	Levinsohn Petrin (2)
Unskilled female	0.34 (.02)	0.47 (.02)
Skilled female	3.10 (.22)	4.98 (.19)
Skilled male	3.34 (.15)	3.68 (.12)
Skilled female-Skilled male (i)	0.93 (.08)	1.35 (.07)
Female-male (ii)	0.72 (.04	0.76 (.02)
Log labor	0.57 (.01)	0.34 (.00)
R-squared	0.72	0.72
N	50,692	252,249

Notes: The equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), industry (38), size (5), age (4), ownership (6), unionization, R&D expenditure per worker, township-and-village enterprise status.

(i) Wage and productivity differentials for skilled women vs. skilled men, and

(ii) skill-weighted wage and productivity differentials for all women vs. all men.

Table A6: Labor input weights for gender-skill groups by firm ownership.

Ownership Sector	Gender-Skill Group	Labor Input Weight	Ownership Sector	Gender-Skill Group	Labor Input Weight
SOE	Unskilled males	1.000	PIE	Unskilled males	1.000
	Unskilled females	0.959		Unskilled females	0.955
	Skilled males	0.992		Skilled males	0.918
	Skilled females	0.945		Skilled females	0.738
COE	Unskilled males	1.000	FIE	Unskilled males	1.000
	Unskilled females	0.951		Unskilled females	1.103
	Skilled males	0.999		Skilled males	0.945
	Skilled females	0.861		Skilled females	0.964

Table A7: Joint production function and earnings equation estimates: Cobb-Douglas production function, robustness checks.

Panel A: Using labor input weights from Table B2	*No weights*			*Weights*		
	Log (Wages) (1)	Log (Value Added) (2)	Column 1 = Column 2 (*p*-value)	Log (Wages.) (3)	Log (Value Added) (4)	Column 3 = Column 4 (*p*-value)
Unskilled female	0.82 (.01)	0.39 (.01)	0.43 (.000)	0.83 (.00)	0.40 (.01)	0.44 (.000)
Skilled female	2.16 (.03)	3.60 (.11)	-1.45 (.000)	2.77 (.04)	4.64 (.14)	-1.87 (.000)
Skilled male	1.82 (.02)	3.38 (.07)	-1.56 (.000)	1.91 (.02)	3.55 (.08)	-1.65 (.000)
Skilled female-Skilled male (i)	1.19 (.02)	1.07 (.04)	0.12 (.005)	1.45 (.03)	1.31 (.05)	0.14 (.000)
Female-male (ii)	0.89 (.01)	0.60 (.01)	0.29 (.000)	0.94 (01)	0.66 (.02)	0.28 (.01)
N	185,245					
Panel B: Proportion of females	*Above median (>0.37)*			*Below median (<0.37)*		
Unskilled female	0.83 (.01)	0.47 (.01)	0.35 (.000)	0.77 (.02)	0.20 (.04)	0.58 (.000)
Skilled female	1.93 (.04)	3.61 (.12)	-1.68 (.000)	2.99 (.05)	4.77 (.19)	-1.79 (.000)
Skilled male	1.88 (.03)	3.42 (.11)	-1.55 (.000)	1.70 (.02)	3.21 (.08)	-1.51 (.000)
Skilled female-Skilled male (i)	1.02 (.02)	1.05 (.05)	-0.03 (.584)	1.75 (.04)	1.49 (.07)	0.26 (.000)
Female-male (ii)	0.81 (.01)	0.53 (.01)	0.28 (.000)	1.07 (.02)	0.79 (.04)	0.28 (.000)
Panel C: Capital intensity	*Above median (>1.75)*			*Below median (<1.75)*		
Unskilled female	0.83 (.01)	0.40 (.02)	0.43 (.000)	0.86 (.03)	0.47 (.01)	0.39 (.000)
Skilled female	2.21 (.04)	3.89 (.16)	-1.68 (.000)	2.02 (.04)	3.25 (.12)	-1.23 (.000)
Skilled male	1.79 (.02)	3.42 (.10)	-1,63 (.000)	1.81 (.03)	3.34 (.08)	-1.53 (.000)
Skilled female-Skilled male (ii)	1.23 (.03)	1.14 (.06)	0.09 (.094)	1.13 (.03)	0.98 (.04)	0.14 (.001)
Female-male (i)	0.93 (.01)	0.70 (.02)	0.23 (.000)	0.89 (.01)	0.57 (.01)	0.32 (.000)

Notes: All equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), industry (38), size (5), age (4), ownership (6), unionization, R&D expenditure per worker, township-and-village enterprise status.

(i) Wage and productivity differentials for skilled women vs. skilled men, and

(ii) skill-weighted wage and productivity differentials for all women vs. all men.

Table A8: Joint production function and earnings equation estimates by industrial sector.

	Metallurgical			*Power*			*Coal*			*Petroleum*		
	Log (Wages) (1)	Log (Value Added) (2)	Column 1 = Column 2 (*p*-value)	Log (Wages) (3)	Log (Value Added) (4)	Column 3 = Column 4 (*p*-value)	Log (Wages) (5)	Log (Value Added) (6)	Column 5 = Column 6 (*p*-value)	Log (Wages) (7)	Log (Value Added) (8)	Column 7 = Column 8 (*p*-value)
Unskilled female	0.76	0.12	0.64	0.59	0.02	0.57	0.48	1.65	-1.17	0.75	0.20	0.55
	(.01)	(.03)	(.000)	(.11)	(.05)	(.000)	(.07)	(.32)	(.000)	(.08)	(.15)	(.000)
Skilled female-skilled male (i)	1.28	1.35	-0.07	0.85	0.41	0.44	1.50	8.54	-7.04	0.76	0.40	0.37
	(.08)	(.18)	(.685)	(.07)	(.07)	(.000)	(.42)	(2.81)	(.010)	(.15)	(.25)	(.159)
Female-male (ii)	0.87	0.44	0.43	0.73	0.25	0.48	0.58	2.73	-2.15	0.78	0.30	0.48
	(.02)	(.04)	(.000)	(.04)	(.04)	(.000)	(.07)	(.37)	(.000)	(.07)	(.13)	(.001)
R-squared	0.84	0.56		0.88	0.81		0.86	0.66		0.90	0.71	
N	27,721			7,460			4,941			2,378		

	Machine Building			*Building Materials*			*Timber*			*Chemical*		
	Log (Wages) (1)	Log (Value Added) (2)	Column 1 = Column 2 (*p*-value)	Log (Wages) (3)	Log (Value Added) (4)	Column 3 = Column 4 (*p*-value)	Log (Wages) (5)	Log (Value Added) (6)	Column 5 = Column 6 (*p*-value)	Log (Wages) (7)	Log (Value Added) (8)	Column 7 = Column 8 (*p*-value)
Unskilled female	0.80	0.23	0.57	0.75	0.43	0.33	0.72	0.32	0.40	0.82	0.53	0.29
	(.01)	(.02)	(.000)	(.01)	(.02)	(.000)	(.02)	(.03)	(.000)	(.12)	(.06)	(.000)
Skilled female-skilled male (i)	1.17	1.34	0.17	1.16	1.14	0.02	1.56	1.02	0.54	1.43	1.53	-0.10
	(.10)	(.05)	(.085)	(.04)	(.06)	(.745)	(.12)	(.17)	(.002)	(.20)	(.34)	(.764)
Female-male (ii)	0.93	0.55	0.38	0.86	0.73	0.13	0.86	0.52	0.34	0.89	0.74	0.15
	(.01)	(.03)	(.000)	(.01)	(.02)	(.000)	(.02)	(.05)	(.000)	(.03)	(.07)	(.035)
R-squared	0.84	0.60		0.85	0.63		0.79	0.54		0.78	0.53	
N	37,451			67,821			18,366			7,429		

Notes: All equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), size (5), age (4), ownership (6), unionization, R&D expenditure per worker, township-and-village enterprise status.

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	(.02)	(.04)	(.000)	(.04)	(.04)	(.000)	(.07)	(.37)	(.000)	(.07)	(.13)	(.001)
R-squared	0.84	0.56		0.88	0.81		0.86	0.66		0.90	0.71	
N	27,721			7,460			4,941			2,378		

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	(.01)	(.02)	(.000)	(.01)	(.02)	(.000)	(.02)	(.03)	(.000)	(.12)	(.06)	(.000)
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	(.01)	(.03)	(.000)	(.01)	(.02)	(.000)	(.02)	(.05)	(.000)	(.03)	(.07)	(.035)
R-squared	0.84	0.60		0.85	0.63		0.79	0.54		0.78	0.53	
N	37,451			67,821			18,366			7,429		

Notes: All equations include a constant term and dummy variables for geographical location (22 provinces, 5 administrative regions and 4 municipalities), size (5), age (4), ownership (6), unionization, R&D expenditure per worker, township-and-village enterprise status.

(i) Wage and productivity differentials for skilled women vs. skilled men, and

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Essay 3: Protecting workers in the workers' state: Labor market tightness and the provision of unemployment insurance in Chinese firms*

1. Introduction

China's ongoing market transition has dramatically reduced the job security of urban industrial employees. Firms have gradually been allowed to hire workers on temporary contracts and to dismiss those already employed. In 1999, "unemployment" was recognized by the government as a formal concept and Unemployment Insurance (UI) was introduced. During the first decade of its existence, the new insurance program was founded on regulations issued by the State Council rather than on a stronger legal basis. A situation thus arose in which enforcement of firm participation was difficult, and evasion easy (Vodopivec and Tong, 2008; Leung, 2003; Jiang and Si, 2009; Sha, 2007). For workers, the combination of weakened job security and, for many, lack of income protection in the case of unemployment, increased economic hardship and the perceived need for precautionary saving[1].

In this paper, I study the effect of labor market tightness, measured by regional[2] unemployment and vacancy rates, on the provision of unemployment insurance in Chinese industrial firms. A simple theoretical model is used to derive the prediction that a tighter labor market leads to increased levels of provision. Intuitively, workers are more attracted to firms that offer UI as part of their compensation package. Thus, in situations where it is difficult to find workers to hire on the labor market, employers have incentives to increase their attractiveness by providing UI.

High quality data are essential to obtain generalizable conclusions about the vast and heterogeneous Chinese economy. For the statistical analysis, I use a uniquely comprehensive panel dataset of Chinese industrial firms for

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[1] An annual survey conducted by China's Academy of Social Sciences in 2001-2005 showed that unemployment was the most commonly cited economic worry among Chinese urban residents (Zhao Huanxin, 2006).

[2] The term region is used for Chinas 22 provinces, 5 Autonomous Regions, and 4 self-governing municipalities.

2001-2005. These data were collected annually by China's National Bureau of Statistics, and include data on firms that together accounted for more than 70% of China's total industrial value added in 2001, and close to 95% in 2005. The comprehensiveness of this dataset, which also includes diverse background information on both the firms and, to some extent, their workforces, enables substantial improvements on the few available studies that have analyzed social insurance program participation in Chinese firms.[3]

The determinants of non-wage compensation have been much debated by labor economists. Even though labor market tightness is the main variable of interest in this study, it is also important to examine alternative explanations of UI provision for three reasons. Firstly, it is important to control for mechanisms that can cloud the statistical evaluation of the main variable of interest. Secondly, examination of alternative explanations allows a critical evaluation of the theoretical model and, thirdly, can indicate fruitful avenues of future research.

Early studies of non-wage compensation provided by workplaces focused on the effects of firm size and unionization (e.g. Rice, 1966; Woodbury, 1983). In the Chinese case, these factors are probably of less importance. In China, insurance is not acquired by firms on the open market. This means that the main mechanism behind the firm size effect, that large firms have a bargaining advantage compared to small ones, is non-existent. Nor does bargaining on the part of workers constitute an important function of China's union. Its main activities are rather to convey party policy and to collaborate with management in smoothing industrial relations (Ge, 2007).

The literature on non-wage compensation has also focused on the role of turnover costs as a motive for supplying benefit plans (Rice, 1966; Mabry, 1973; Long and Scott, 1982; Woodbury, 1983). To avoid the cost of replacing labor, especially for workers with high skills, firms offer benefit plans that cannot easily be moved between workplaces.

Another factor studied in the early literature, related to worker demand for benefits, was the effect of tax exemption of benefit plans (e.g. Long and Scott, 1982). Under a progressive income tax system, the higher the worker's wage, the greater the gain would be for him or her to receive their work compensation in the form of tax exempt benefits rather than in wages. From a political science perspective, Mares (1997; 2003) argued for the role of labor market risks for worker demand for unemployment insurance.

In a transition economy, a firm's ownership type is often an important driver of its behavior, owing to the close links between ownership type and

3 Nielsen and Smyth (2007) use audit data for 5,212 firms in 2002 and 5,580 firms in 2003, all located in the city of Shanghai and selected by the city's Bureau of Labor and Social Security; and in a Ph.D. thesis in sociology, Lu (2007) analyzes data for 422 State-owned firms in a yearly 1990-1999 panel.

the legal and financial situation of the firm[4]. In China, the private sector's subordinate status relative to the "socialist public economy" is defined in the constitution. In practice, this has limited private firms' entry into industrial sectors, and their access to capital. Besides the favorable treatment of state-owned enterprises (SOEs) compared to private firms, foreign owned firms have also held a somewhat advantageous position (Huang, 2003). In addition to effects on the "ability to pay", ownership type is likely to influence the level of scrutiny by the authorities[5].

The results presented in this paper support the prediction that firms operating in tighter labor markets are more likely to provide their workers with unemployment insurance, but the effect is quantitatively small. Inspection of the control variables included in the empirical analysis also provides some interesting preliminary findings. Unlike labor market conditions, firm and workforce characteristics, such as ownership, wages and firm size, appear to explain a substantial part of the variation in firms' insurance provision.

Improvement of our understanding of these determinants has an importance that reaches beyond the academic literature. Improvement of workers' income stability would lead not only to improved welfare, but would also probably decrease personal savings[6]. Increased household consumption would, in turn, allow for a rebalancing of China's sources of growth away from foreign demand and toward domestic consumption. It is also noteworthy that the combined firm payments required by China's social insurance system are substantially higher than in other countries, amounting to 37-53% of the unit labor cost. Insights into the pattern of participation, and the mechanisms that affect firm choice, are thus potentially crucial for analyzing the Chinese business environment and the competitiveness of China-based firms in the international marketplace.

2. The evolution of China's unemployment insurance

Shortly after the proclamation of the People's Republic of China in 1949, an ambitious set of entitlements for the urban workforce was formulated. Since

4 Poncet et al. (2009) provide micro data evidence that the private sector faces harder capital constraints than the state-controlled sector and Sourafel et al. (2008) add evidence for a bias in production subsidies toward SOEs.

5 The firm audit datasets from Shanghai's Bureau of Labor and Social Security (BOLSS) used by Nielsen and Smyth (2008) and Nyland et al. (2005) show substantial changes in the year-on-year ownership distribution. Citing interviews with the BOLSS, Nielsen and Smyth (2008) explain the sharp changes in terms of the BOLSS' perception of low compliance among certain groups.

6 Private savings as a share of disposable income has consistently risen since the second half of the 1990s, reaching 28% in 2008 (Prasad, 2009). From a macroeconomic perspective, these increasing savings have been a main contributor to the declining share of consumption in China's GDP, alongside the declining share of household disposable income as a share of GDP, and a low level of government consumption (Lardy, 2007).

firms were not allowed to fire their workers, the so-called the Labor Insurance Scheme (LIS) did not include provisions for income protection in the case of unemployment. However, after Mao's death and the start of China's gradual reform of the planned labor market, things began to change.

A restructuring of the state-controlled sector started in the 1980s and necessitated the transfer of a large number of workers between jobs. To ensure that the basic needs of these "temporary jobless" persons were met, the *Interim Provisions on Unemployment Insurance of Staff of State Enterprises*, in short, the "Interim Provisions", were put in place by the State Council in 1986. Firms were asked to contribute 1% of their total payroll to the system in order to provide for the protection of four categories of staff[7]. The benefit level was set at between 60-75% of the average monthly wage, but could vary with the length of service and other basic criteria. In the second year of temporary joblessness, the benefit rate was reduced to 50%. Rather than to provide income insurance, the main purpose of the system was to assist workers in job seeking, job training and referrals.

As market reform continued, the Interim Provisions became inadequate in the face of the growing number of unemployed. In 1993, the provisions were amended by a regulation package called the *Regulations on Unemployment Insurance for Staff and Workers of State-Owned Enterprises*. The amendments extended the program's coverage to three more SOE staff categories[8]. Furthermore, the link between the benefit level and individual wages was removed. Instead, benefits became a function of the total amount of social relief payments stipulated by the local institutions of civil affairs. Finally, with the reform, workers who refused job offers twice without proper reasons were excluded from benefits (Vodopivec and Tong, 2008).

The job-guarantee under the planned economy eroded during the early 1990s as employers were gradually given increased autonomy over workforce decisions. This new autonomy was extensively practiced in the continued reconstruction of the state-owned sector during the latter half of the 1990s. It is important to note that workers laid-off during this period did not become unemployed in the sense of being sent out to look for work in a flexible labor market. Instead, they remained with their former state-owned work units as so-called job-waiters (x*ia gang*). SOEs were required to establish Re-employment Service Centers to cater for their Xia gang workers. The number of persons registered at the centers peaked at 7 million in 1991, but

7 These were: (i) staff of an enterprise adjudicated bankrupt; (ii) staff discharged in the statutory reorganization of an enterprise on the verge of bankruptcy; (iii) staff discharged on the rescission of a labor contract or on the dissolution of an enterprise; and (iv) staff dismissed by an enterprise (Lee, 2000).

8 These were: (i) staff and workers of enterprises closed or dissolved in accordance with relevant regulations of the State; (ii) staff and workers laid off in periods during which the enterprises ceased production in order to be streamlined in accordance with relevant regulations of the state; and (iii) workers who had their employment contracts terminated or canceled.

had fallen to 200,000 by 2005 (Vodopivec and Tong, 2008). The key reason for the large reduction was the so-called *binggui* policy, which transferred the job-waiters from their former employers to the present day unemployment insurance system.[9]

A major limitation of the protection from joblessness in China's early transition period was the limited scope of the regulations in place. A growing number of workers outside of the state sector lacked any protection, and even within the state sector, coverage did not encompass all staff. This changed in 1999, when China formally acknowledged the concept of "unemployment"[10] and set up an insurance system similar to those of western nations. With the adoption of the *Regulations on Unemployment Insurance* firms in all ownership sectors were required to provide UI coverage to their employees. However, as under the LIS, firms in rural locations, and workers without permanent labor contracts, were largely excluded.

The basic principles of the UI program's funding and benefits are the following. Benefits are not primarily calculated based on earnings, but instead on the total period with continuously paid contributions. According to this principle, 1 to 5 years of contributions merits 12 months of benefits, 5 to 10 years gives 18 months, and more than ten years gives the maximum 24 months. A worker becomes eligible for benefits under the conditions that: i) his or her enterprise has continuously paid premiums for at least one year[11], ii) the termination was involuntary, and iii) he or she is registered as unemployed and willing to be re-employed. The level of the benefits is set by the authorities of provinces, autonomous regions and municipalities, with the requirement that it is above the minimum living standard of urban residents[12].

Payments are made by both the employer and the employee, but the employer is responsible for depositing the funds. With minor regional deviations, the contribution rate is 2% of total payroll for firms, and 1% of wages for workers. These payments are not subject to profit tax in the case of the firm's share, or income tax in the case of worker's (Vodopivec and Tong, 2008).

Two important aspects of the UI regulations promulgated in 1999 are that they were vague and not legally binding. As a consequence, authorities at the

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9 This transition was first completed in Beijing, Tianjin, Shanghai, Jiangsu, Zhejiang, Guangdong, Fujian, Shandong and Liaoning. In 2005, a total of 20 provinces had completed the binggui (Vodopivec and Tong, 2008).

10 Following the ILO definition, persons were defined as being unemployed if they became jobless: i) as a consequence of enterprise bankruptcies; ii) by having their contracts terminated or canceled; iii) by being dismissed; iv) by resigning; or v) after graduating and failing to find first-time employment.

11 Note that the firm is responsible for depositing both the firm's share and the worker's share of the contributions.

12 As such, the income protection of the program is low, representing a replacement rate of 14.7% in 2005 (Vodopivec and Tong, 2008).

regional level had substantial freedom to influence the specific provisions of their region, but at the same time they lacked legal backup for enforcing firm compliance. In essence, firms could choose between making their required payments or facing small repercussions of evasion. Not surprisingly, many evaded and the descriptive statistics provided by Vodopivec and Tong (2008: Table 1) actually document a decrease in the proportion of covered workers from 45% in 2000 to 39% in 2005.

After several years of deliberations, China's social insurance programs, including the UI program, were set on a more firm legal footing with the passing of a Social Insurance Law in late 2010.

3. Studies of social insurance provision in Chinese firms

There are only a small number of published papers and unpublished reports that deal with the issue of firm provision of social insurances in China. Lu (2007) analyzes data for 422 state-owned firms using a yearly panel from 1990 to 1999. He finds that market competition and exposure to other successful and complying firms were important drivers of firms' social security spending. Also, he finds higher spending in firms with funding from the central government rather than from local governments[13], in firms where the CEO had party membership, and in larger, older, and more profitable firms.

Nyland et al. (2005) use data for about 2,000 firms audited by Shanghai's Bureau of Labor and Social Security in 2001 to investigate compliance by firm size, ownership and industrial sector. They find that larger firms were less likely to meet their prescribed social insurance liabilities. Firms operating in the banking, construction and real estate sectors were also more likely to comply. In terms of ownership, less compliance was found for foreign-invested, HKMT (Hong Kong, Macao, Taiwan)-invested, and shareholding firms. In interviews, eight managers mainly emphasized the economic rationale of evasion, citing the low probability of being audited combined with the weak enforcement mechanism once transgressions were exposed[14].

Nielsen and Smyth (2007) use similar Shanghai audit data for 2002 and 2003, to investigate the extent to which firms' costs of social insurance payments are shifted over to workers through the lowering of wages. This inci-

[13] Regarding this particular finding, Lu (2007) argues that firms with funding from the central government are less sensitive to market uncertainties and **are** more likely to be bailed out. Meanwhile, firms that rely more on retained earnings would be more likely to have a larger degree of managerial autonomy and be more concerned about raising firm costs.

[14] Non-compliant firms were identified by means of annually conducted random audits of about 2-5% of the total population of firms. If found to be in arrears, firms were first given a 15 day period to make up their payments. After this period, interest of 0.02% was added to the amount in arrears, and after 10 months, the authorities could mortgage firm property to cover the debt. A hotline for employee complaints was also available, but judged as a poor enforcement method by the interviewed managers.

dence is estimated to be 9.1% in 2002 and 33.8% in 2003, which are both low compared to standard results for other countries. The descriptive statistics also cast some doubt on the authors' claim that the BOLSS dataset constitutes a representative sample of Shanghai's firm population. The proportion of firms meeting their payment obligations tripled from 18% in 2002 to 65% in 2003, and the ownership distribution shifted dramatically between the two years[15].

Returning to the participation decision, unpublished studies drawing on smaller firm samples also contribute some insight. Jiang and Si (2008) survey 101 private firms in five cities in 2007. In an informal analysis of their results, they notice a strong positive correlation between participation and firm size. This effect is conjectured to stem from several factors correlated with firm size, in particular the need to attract talented workers and to protect the firm's image at home and abroad. In the firms' own opinions, cost considerations were listed as the main driver of their compliance decision. Similar results were obtained by Sha (2007) in a survey of 100 present or former state-owned firms in the coastal province of Jiangsu and the inland province of Sichuan in 1999. Loss-making firms reported an inability to afford to make payments, and privately owned firms in particular perceived social security spending cuts as a useful cost-saving measure. In contrast to private firms, the above-average participation rate among foreign firms was, according to managers, mainly related to their desire to comply with government requests and to alleviate uncertainty among employees.

Both Sha (2007) and Jiang and Si (2008) identify effects due to worker demand factors such as unionization and the proportion of workers with permanent contracts. In particular for private firms, migrant workers were reported by managers to prefer wage compensation to insurance coverage owing to the weaker insurance rights implied by their migrant status. Emphasizing bargaining power, Sha (2007) finds that intense competition on the labor market coupled with a high turnover rate is associated with the lower bargaining power of the mainly temporary workers in the private sector and thus lowers social security spending among private firms.

4. A simple model of the firm's UI choice

Consider a firm operating in a fictional labor market where local labor market conditions influence the speed of recruitment. Employment, N_i, in firm i changes according to

15 The share of SOEs in the sample declined from 39% in 2002 to 12% in 2003 and the share of private firms dropped from 10% to 4%, while the share of foreign-invested firms rose from 33% to 43%.

$$\partial N_i / \partial t = H_i - \phi N_i \ , \tag{1}$$

where H_i is the flow of hires and ϕ the exogenous separation rate. The hiring function takes the form

$$H_i = qV_i \, , \tag{2}$$

where q is the rate at which vacancies are filled and V_i is the number of vacancies. The rate at which vacancies are filled depends on local labor market characteristics, such as the local vacancy rate (v) and the local unemployment rate (u). It also depends on compensation to workers, such as wages and social insurance benefits.

Assume that the firm is a wage taker but that it can influence recruitment by provision of UI. Suppose that the firm's recruitment function is given by

$$q = a(b)\theta^{-\eta} \ , \tag{3}$$

where $\theta = v/u$ is a measure of labor market tightness, $a(b)$ captures a positive relationship between recruitment intensity (a) and UI benefits (b) with $a'(b) > 0$ and $a''(b) < 0$, and η is a positive constant. The tighter the local labor market, the more difficult it is for the firm to fill a vacancy, but the more generous benefits the firm offers, the easier it is to fill the vacancy.

The firm's profit function is

$$\pi_i = AN_i^{\alpha} - wN_i - T(b)N_i - cV_i \quad \alpha < 1 \, , \tag{4}$$

where w is the wage rate, T is a tax function that is increasing in UI provision, T '(b) > 0 and c is the cost of recruitment. The firm ignores discounting, takes the wage as given and chooses N_i and V_i to maximize steady-state profits, recognizing the hiring constraint $a\theta^{-\eta}V_i = \phi N_i$. Thus,

$$\pi_i = AN_i^{\alpha} - wN_i - \left[T(b) + \frac{\phi c \theta^{\eta}}{a(b)} \right] N_i \ , \tag{5}$$

and the first-order conditions for a maximum are

$$N_i : \quad \alpha A N_i^{\alpha - 1} - w - T(b) - \frac{\phi c \theta^{\eta}}{a(b)} = 0 \, , \tag{6}$$

$$b: \quad -T'(b) + \frac{\phi c \theta^{\eta} a'(b)}{[a(b)]^2} = 0 , \tag{7}$$

which determine N_i and b as functions of the wage and labor market tightness. Provided that an interior maximum exists, it is straightforward to establish that $\partial b / \partial \theta > 0$. The marginal gain to the firm of (increased) UI provision is higher the tighter the labor market is.

To understand this implication, note that the cost of a vacancy is

$$cV_i = \frac{N_i \phi c \theta^{\eta}}{a(b)} , \tag{8}$$

which means that the tighter the labor market is, the greater is the cost-reducing effect of an increase in *b*.

The result is perhaps surprising since one might have expected that slack labor markets with high unemployment would trigger an increase in worker demand for UI. Such mechanisms are perhaps relevant, but require channels though which demand for UI carries over to actual UI provision.

5. Data and descriptive statistics

5.1 Panel data for Chinese firms

The firm panel data used in this study come from the annual firm survey conducted by China's National Bureau of Statistics. It covers all state-owned enterprises and all other enterprises with annual sales exceeding 5 million RMB. In total, the number of observations ranges from 164,500 in 2000 to 255,546 in 2005. These firms together account for a large fraction of China's total yearly industrial GDP, roughly 70% in 2000 and 95% in 2005[16].

Over time, firms in the panel may change ID codes because of restructuring, mergers or acquisitions. To correct for these changes, I use background information on firms, such as location, industry etc. to expand the panel. To reduce misreported firm information, I follow Jefferson et al. (2008) and drop firms with fewer than eight employees and firms not in operation. The industrial sector classifier is corrected for the code change in 2003[17] and

16 Calculated by dividing the total value added of the firms in the dataset by the yearly industrial GDP in China's Statistical Yearbook.

17 In this year, the classification system GB/T 4754-2002 replaced the GB/T 4754-1994 system. A basic concordance table shows only slight changes of products between new and old 2-digit divisions on the 3-digit level. On the 2-digit level, the following changes are made:

deflation of the cost series into 2001 prices is done using the Ex-Factory Price Indices available in China's Statistical Yearbook[18].

Table 1: Summary statistics, 2001-2005.

	2001	2002	2003	2004	2005
State-owned (share)	0.23	0.18	0.14	0.09	0.07
Collective-owned (share)	0.25	0.21	0.17	0.10	0.08
Private domestic (share)	0.22	0.28	0.35	0.42	0.46
Foreign-invested (share)	0.07	0.08	0.19	0.10	0.11
HKMT-invested (share)	0.11	0.11	0.11	0.10	0.10
Average monthly wage (¥)	848	906	970	1,125	1,223
Employment	321	312	298	244	258
Profit per worker (¥)	11,017	13,581	15,586	17,690	20,102
Rural firm (share)	0.52	0.46	0 .41	0.17	0.19
Market share	0.029	0.027	0.025	0.018	0.018
Subsidies per worker (¥)	643	712	-	977	1,029
Number of observations	162,436	169,957	190,252	267,289	255,163

Notes: Information on firm subsidies is not available for year 2003. "Mixed" ownership not reported.

Summary statistics for the remaining yearly samples are reported in Table 1. In the top five rows we can trace the ownership diversification of China's industrial sector in 2001-2005 as seen in the declining share of state- and collective-owned firms, and the rapid growth of the private domestic sector.[19] Simultaneously, the share of "rural" firms dropped from 38% in 2000 to 9% in 2005.[20] In the bottom row of the table, we see that the sample size peaks in 2004. In this year, a firm census was conducted and more firms were reached by the authorities. A thorough comparison of the descriptive statistics for the additional firms sampled in the census year with the full 2003 and 2005 samples does not however reveal any notable deviations.

first, the division "wage recycling" is dropped prior to 2003. Second, the division "weapons and ammunition manufacturing" is added to the "special equipment manufacturing" category after 2003. Third, the division "Bamboo and Lumber" is added to the "Timber" category for 2003 onward.

[18]Firms are grouped into 14 sectors: Metallurgical, Power, Coal, Petroleum, Machine Building, Timber, Chemical, Food, Textile, Tailoring, Paper, Culture and Education, and Leather.

[19] Ownership categories are created by aggregating the 23 registration types into six broader categories, closely tracking the formal classification system currently used for reporting data on all industrial firms in China's Statistical Yearbook (see Table A1 for details). The somewhat unusual term collective-owned is used for urban and rural firms with investment from collectives, and some firms registered with industrial and commercial administration agencies as collective units. In the latter case, funds are pulled together by individuals who voluntarily give up their right of ownership (China Statistical Yearbook 2006, chapter 13). Collective firms are often managed by representatives from provincial or local government bodies.

[20] A firm is considered to be "rural" if it is under the administrative control of a county, small town, street, village, resident committee or village committee.

5.2 Unemployment insurance provision

The survey data provide a firm's combined annual expenditure on unemployment insurance (UI) and the Labor Insurance Scheme (LIS). Although there has been a requirement for state-controlled firms to transfer from the LIS to the separate social insurance programs since 1999, some SOEs retained some or all of their LIS provision during the period of study. Hence we face the potential measurement problem of misinterpreting ongoing participation in the LIS with participation in the UI program.

I use an indicator variable for UI provision, rather than a continuous measure such as the insurance cost per worker, since this has the advantage of minimizing measurement error due to over-reporting of provision. Over-reporting of social insurance contributions, as well as under-reporting of wages, lowers a firm's tax burden (Banister, 2005) [21]. There are then a number of approaches to the measurement problem noted above. Firstly, we can simply assume that non-zero provision in the combined variable (LIS+UI) implies some provision of UI. Secondly, we can exclude SOEs since only these firms participated in the LIS. Thirdly, we can exclude firms that report spending more than the UI maximum of 2% of payroll.

Figure 1 shows the average provision rate over time of the full sample and for the four ownership types, using the dichotomous measure of provision. Although the full sample average remained fairly constant, at around 36-40% over the 2001-2005 period[22], there is more variation within the different ownership types. Notably, there are higher-than-average rates of provision among firms with state or mixed ownership, and among firms with non-mainland investment, i.e. from Hong Kong, Macao or Taiwan (HKMT), or from foreign entities. Collective-owned and private domestic firms have lower rates of provision over the period.

Figures 2 and 3 show the distribution of average UI provision rates across regions and industries over time. Calculation of average provision rates by region shows that above-average rates are generally found in more developed localities, such as Beijing, Shanghai and Fujian, while poorer regions like Tibet or Gansu have rates that are below average. The highest average provision rates by 2-digit industry code are for heavy industrial sectors, such

[21] Evidence of widespread fraudulent tax reporting was also found in the World Bank's World Business Environment Survey (WBES), carried out at 100 firms in each of 80 Chinese counties in 2000. In the survey, only 11.9% of the Chinese firms thought that 100% of the other firms in their industry gave accurate reports of their incomes to authorities.

[22] It is important to note that changes in the average provision rate over time come only from the net number of exits and entries into the UI program. If we take an "entry" to be a firm having zero costs in year t-1 but positive costs in year t, and the opposite for "exits", then there are increasing trends in both the number of entries and the number of exits over time. The number of entries increases from 13,240 in 2001 to 33,148 in 2005; while the corresponding growth in the number of exits is from 13,684 to 21,450.

as the petroleum industry, and for utilities. Meanwhile, firms in light industrial sectors appear more likely to evade.

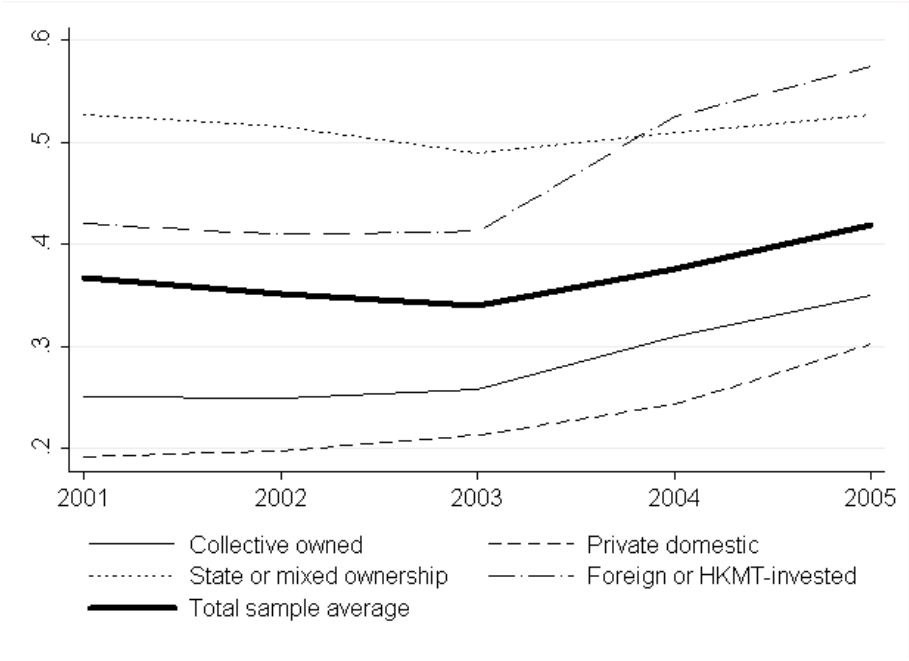

Figure 1: Share of UI provision, full sample and by ownership type, 2001-2005.

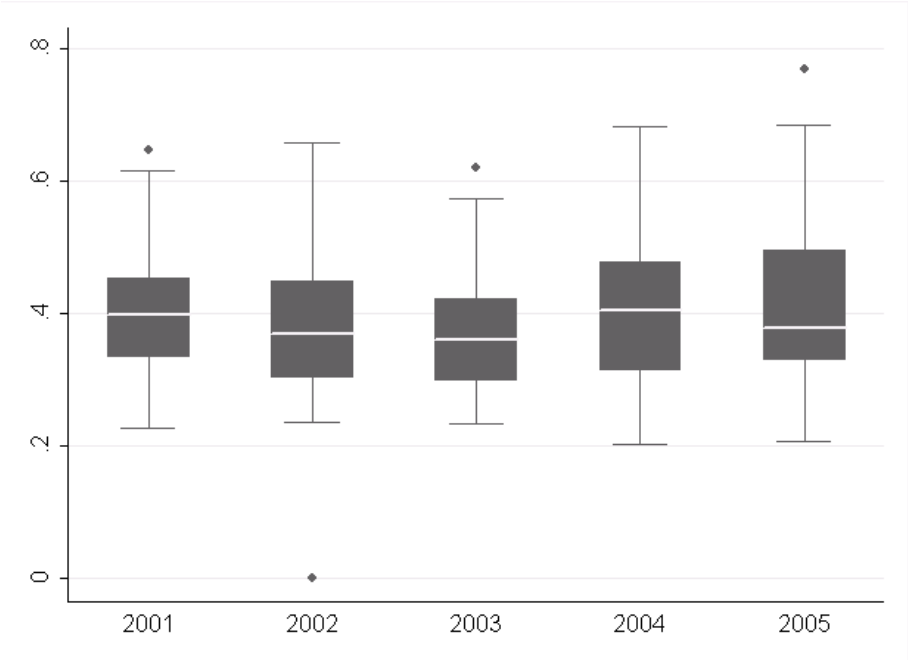

Figure 2: Distribution of regional average rates of UI provision, 2001-2005.

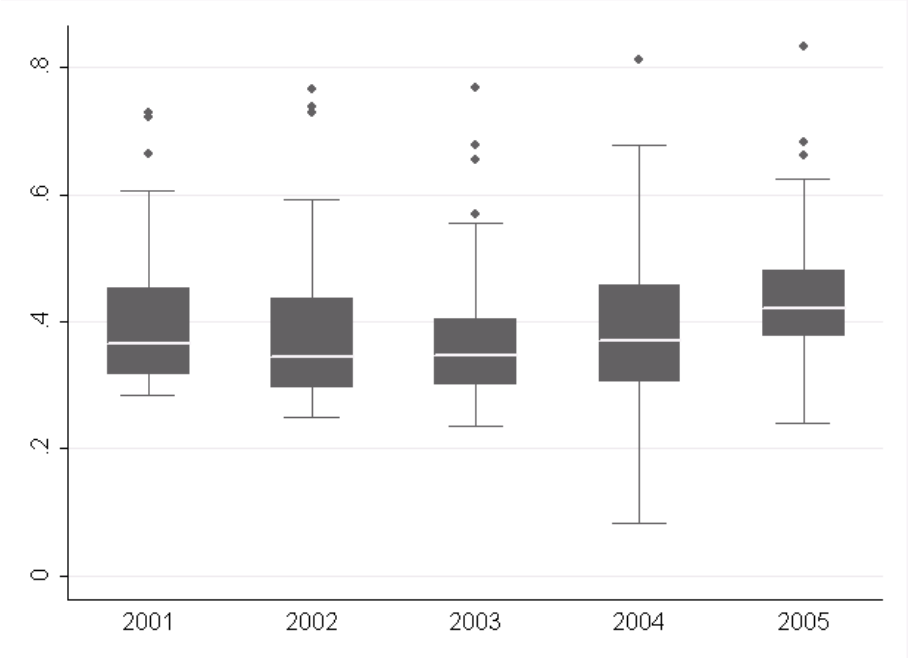

Figure 3: Distribution of 2-digit industry average rates of UI provision, 2001-2005.

5.3 Labor market tightness

I use the official measures of regional unemployment and vacancy rate. These are based on the registration of unemployed persons and vacant jobs with regional unemployment offices. A comparison of these headcount measures with regional urban labor force numbers yields the unemployment and vacancy rates reported in China's Statistical Yearbook. The reliability of China's official measure of the unemployment rate has been the subject of some debate. Several studies have tried to calculate the "true" rate of unemployment (see e.g. Giles et al. 2005). The usefulness of such alternative measures for the purpose of this study is limited, mainly due to the lack of calculations at the regional level, but also because of the lack of consensus on the "correct" calculations.

The main criticism of the official unemployment measure is the exclusion of the laid-off SOE job-waiters (*Xia gang*). However, simply adding the Xia gang workers to the official unemployment headcount is not advisable for the simple reason that the Xia gang workers often have jobs.[23] Instead of pursuing this strategy, I therefore use information on the regional variation in the *binggui* reform, whereby the job-waiters were pooled with the official unemployment statistics, as a way to reduce measurement error in the unemployment variable. As noted above, 20 regions had completed this transition by 2005.

A second serious criticism of the official unemployment count is the exclusion of so-called "migrant" workers who have moved to urban areas but lack the official right to either work or be unemployed there.[24] At least according to the official view, excluded migrant labor only marginally affects the unemployment statistic. This is because migrants self-select into employment and return to their rural hometowns when the urban labor market becomes tighter (Fox and Zhao, 2002).

Figures 4 and 5 show the distributions of unemployment and vacancy rates across regions over time. Figure 6 shows the distribution of the labor market tightness measure (the ratio between the vacancy rate and the unemployment rate) across regions over time.

[23] Such "hidden employment" may also exist in the unemployment numbers, in particular because of low UI benefit levels and weak enforcement of eligibility criteria.

[24] Using data from the second round of the China Urban Labor Survey (CULS 2005), Du et al. (2006) find that 32.5% of native urban residents and 84.3% of migrants did not have formal labor contracts in 2005.

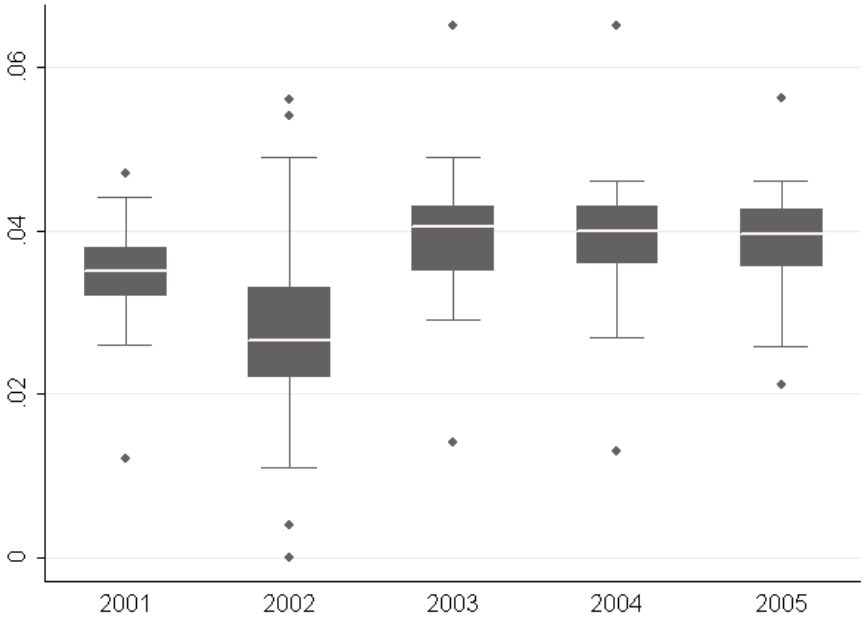

Figure 4: Distribution of regional unemployment rates, 2001-2005.

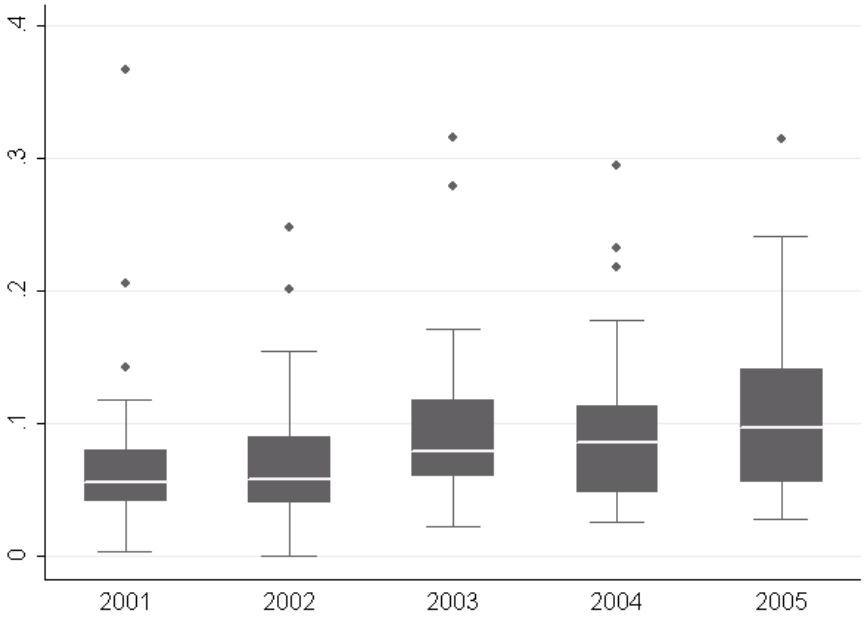

Figure 5: Distribution of regional vacancy rates, 2001-2005.

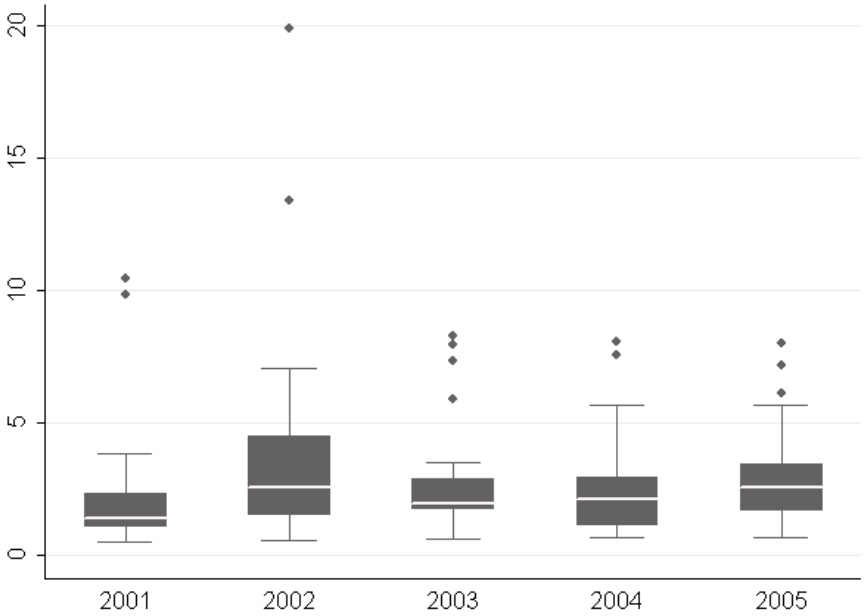

Figure 6: Distribution of regional levels of labor market tightness (vacancy rate divided by unemployment rate), 2001-2005.

6. Empirical specification

Predictions from the theoretical model are tested using the following specification

$$UI_{i,t} = \beta_0 + \beta_1 \theta_{k,t} + \boldsymbol{\beta}_2 \mathbf{S}'_{i,t} + \boldsymbol{\beta}_3 \mathbf{Z}'_{c,t} + \boldsymbol{\beta}_4 \mathbf{P}'_{k,t} + \varepsilon_{i,t}, \tag{10}$$

where $UI_{i,t}$ is a dichotomous variable for firm *i*'s UI provision in period t. Our main variable of interest, $\theta_{k,t}$, is the tightness measure in region k over time.

A host of control variables is included in the empirical specification to take account of factors influencing a firm's UI provision other than the situation in the regional labor market.[25] First, the firm-level controls are contained in the vector $\mathbf{S}'_{i,t}$. The average monthly wage per worker in the firm is included to capture turnover costs (e.g. Rice, 1966). I also include a categorical variable for the modal tax bracket (1-6) that the average worker of the firm belongs to. The expected sign of this variable is uncertain. On the other

[25] Additional information about the specification of the control variables is summarized in Appendix, Section A1.

hand, Chinese high-wage earners may prefer to receive wages instead of UI, regardless of the tax exemption, because of the program's low, flat compensation rate (e.g. Long and Scott, 1982).

Firm size and market share could both be expected to affect the likelihood of participation since they shape the firm's cost profile. Large firms with large market shares have a greater capacity to pass payment obligation costs on to customers via a higher product price, while small firms have to face the full insurance cost (Mares, 1997; 2003). To further control for the firms' finances, I include measures of the profit per worker and the amount of state subsidies per worker the firm receives.

Several firm-level measures are included to capturing the firm's institutional setting. Ownership dummies are used to group firms into six sectors. As noted in the introduction, ownership type can determine the firm's access to capital and the level of monitoring scrutiny from social insurance authorities. Owing to these factors, we expect higher levels of provision for state-owned firms and for foreign-invested firms than for private domestic firms; in accordance with the institutional "pecking order" proposed by Huang (2003). Perhaps more important for the research question at hand, firm ownership can also be expected to be correlated with the way firms interact with workers on the labor market. Firms in private ownership, both domestic and foreign, can be expected to more often be in competition with other firms in the hiring of workers than can firms in the state-controlled sectors, and may thus be more responsive to labor market conditions in determining whether or not to provide UI. Finally, firms under different forms of ownership may be expected to systematically hire different types of labor. Private sector firms are known to be more likely to hire younger workers and migrant workers; the latter are expected to have lower UI demand.

To further capture the institutional setting, I control for the rural status of firms, and for firm age. Lower participation rates can be expected for rural firms because they often operate under less restrictive social insurance regulations. Lower participation rates can also be expected for older firms because they are less likely than newly established firms to hire workers on the open labor market.

Some additional workforce characteristics are available for the census year 2004. They measure firm unionization, the share of female workers, and the human capital distribution of the firm's workforce. All three variables are included to control for the potential effect of worker demand, and, in the case of skilled workers, for the cost of labor turnover.

A second vector of controls $\mathbf{Z}'_{c,t}$ is available for 284 cities z, providing variable values for over 85% of the firms in the dataset. Two of these controls are included in an attempt to approximate the level of enforcement by local authorities. They measure the distances between each city and the closest of China's three large ports, and the distance to the closest of China's 15 largest cities. Geographical location determines the potential of the locality

to attract investment. To avoid scaring off potential investors with high social security requirements, authorities in less advantageous locations are expected to be more lenient in enforcing compliance with social security regulations.[26]

A third city-level control variable uses data from 2004 and 2005 to approximate the average total social security cost faced by firms. The importance of this control variable is in the fact that even though the UI program in itself is not expensive for the firm as a share of wages, participation in the UI program makes it more difficult for the firm to avoid the costs of other insurance schemes For each city, I calculate the total insurance cost as a share of wages for firms that participate in all major schemes, i.e. those for pensions, medical insurance, and housing accumulation funds. The cost levels of all these programs can, and do, vary at the prefecture-city level.

Unlike the other social insurance programs, the contribution level of the UI program is determined at the regional level. Data for these cost levels are included in the vector $\mathbf{P}'_{k,t}$.[27] At the regional level, I also control for the share of state-owned firms in total employment. I expect this to be negatively correlated with the UI provision rate owing to the legacy of unmet obligations in the state-owned sector, which has reduced public confidence in the social insurances as well as the expected benefits among firms and workers (e.g. Jackson et al., 2009).

Since firm-level data are used, the analysis may be subject to bias due to unobserved worker heterogeneity. Although the dataset contains information on the human capital distribution and gender composition of firm workforces, information on the workers' age and labor contract type is lacking. This could be problematic if some firms systematically seek to hire workers that are, for example, young and migrant, since UI demand is expected to be lower among these groups. Although such unobserved factors may influence a firm's provision of UI, the included measures of ownership type, industry, location, age of firm and workers' education level can be expected to capture some of these effects.

[26] Leniency or stringency in monitoring may be an important link between the institutional setting of a firm and its compliance behaviors, as suggested by McGillivray (2001). Anecdotal evidence in China is that local authorities may not want to "put a damper" on the development of the private economy by imposing harsh social security regulations (Xinhua 2006-12-06). In disadvantaged locations, local officials may also believe that it is better for an employer to provide jobs without social security than no jobs at all.

[27] The cost level rarely deviates from the state's guideline of 2% of the worker's average annual wage. Exceptions are: Beijing, 1.5% in all years; Ningxia: 1% in 2001; Tibet 1% for SOEs only in 2001 and 2002, 2% for SOEs in 2003; Hebei 1% in 2000; Hubei 1% in 2001 and 2003; Hunan 1% for SOEs only in 2001; Jilin 1.5-2% except for 1% for SOEs in 2001 and 2003; Shaanxi 1% for SOEs only in 2001 and 2002, 2.5% for all firms since August 2003; Shanxi 1% in 2001-2004; Sichuan 1% in 2001, but rates down to 0.6% allowed in cities; Yunan 1% in 2001-2003 (as indicated on province homepages).

7. Results

7.1 Labor market tightness and firm provision of UI

Baseline results from estimating a model for UI by OLS are given in Table 2. All standard errors are clustered at the regional level to take into account the fact that unemployment and vacancy rates are measured at this level. All estimations also include year fixed effects to account for time trends in the average level of UI provision. They also include fixed effects at the 2-digit industry level.

The theoretical model predicts that a tighter labor market (lower unemployment and more vacancies) should increase the likelihood of firm participation in the UI program.

I first model the UI provision rate as linear in the unemployment and vacancy rates. The OLS parameter estimates for this model are given in Table 2, Column (1). The estimated parameters take on the expected signs, but of the two estimated parameters, only the one for the vacancy rate is statistically significant. Since these variables are measured on a scale from 0 to 1, the parameter estimate for the vacancy rate implies that an increase of the vacancy rate by one percentage point leads to a 0.6 percentage point increase in the likelihood of UI provision.

Turning to the labor tightness measure, I examine the robustness of its parameter estimate by estimating Equation (10) without the control variables. The parameter estimate is given in Table 2, Column (2). It has the expected positive sign and enters at the 1% significance level.

The results of estimating the full baseline model with controls are given in Table 2, Column (3) (note that only select controls are reported in the table). When the controls are included the parameter estimate for labor market tightness is halved, although it remains highly significant and with a positive sign.

I next address some concerns regarding measurement error in the dependent and the main independent variables. To avoid erroneously counting firms that participate in the LIS rather than in the UI program, in Column (4) results are given for the case where state-owned firms are removed, and in Column (5) results are given for the case where firms that report spending more than 2% of their total wage bill on unemployment insurance provision are removed. Examining the parameter estimate for the labor market tightness variable, we see that it is not significantly altered by these sample adjustments.

Turning instead to the possible error in the labor market tightness measure, I use information on which regions had completed the *binggui* policy by 2005. Exclusion of all but these regions should reduce any measurement error stemming from uncounted job-waiters in the regional labor markets.

The estimation results in Table 2, Column (6) show that the parameter estimate for labor market tightness is insensitive to this sample change.

So far, the estimations have included year and industry fixed effects only. I now include regional fixed effects (Table A3, Column 1) and firm fixed effects (Table A3, Column 2). In these estimations, labor market tightness drops out of the model. This is probably because the inclusion of year, industry, and either regional or firm fixed effects removes too much of the variation in the independent variable to enable identification of its parameter. The baseline result is hence identified from differences between regional labor markets rather than from trend-deviations within each region over time. We also know that the results are not driven by differences in the industry composition between regions.

I evaluate the sensitivity of the baseline parameter estimates to changing the estimation method from OLS to probit. A comparison of the baseline estimates (Table 2, Column 3) to the marginal effects from the probit estimation (Table A3, Column 3) shows that they are very close.

The results from some additional robustness checks are found in Table 3. In these regressions I use the balanced panel of firms, which reduces the sample size to 52,728 valid observations per year. Descriptive statistics for the balanced panel are reported in Table A2. The parameter estimates for the baseline specification on the balanced panel are shown in Table 3, Column (1). We can see that the labor market tightness parameter estimate is somewhat smaller than previously, but is still highly significant. Next, an attempt is made to control for some aspects of workforce heterogeneity. Worker education and firm unionization were available for 2004 only, and the share of females for 2004 and 2005 only. These are assumed to be constant over the full 2001-2005 period. The estimate of the effect of labor market tightness on UI provision is robust to the inclusion of these controls.

Up to this point, I have estimated effects on participation using a dependent variable that is based on both entries and exits from the UI program. As noted above, this could include firms that dropped out of the LIS without replacing it with the UI program. To take this issue into consideration, I re-estimate the model for entries only. In the two rightmost Columns of Table 3, the dependent variable is a dummy capturing whether a firm entered the system in the subsequent year. The parameter estimate for the effect of labor market tightness remains positive, significant and small.

Table 2: OLS estimates of the effects of labor market tightness on firm provision of unemployment insurance. The dependent variable is 1 for firms reporting some non-zero insurance payment in a given year, otherwise 0. Unbalanced panel data from 2001-2005.

	Full sample (1)	Full sample (2)	Full sample (3)	Excluding SOEs (4)	Excluding high UI payments (5)	Excl. non-binggui regions (6)
Unemployment Rate	-0.819					
	(1.141)					
Vacancy Rate	0.577**					
	(0.269)					
Labor market tightness		0.013***	0.006**	0.008**	0.008***	0.008**
		(0.004)	(0.002)	(0.003)	(0.002)	(0.003)
Average wage			0.087***	0.083***	0.085***	0.080***
			(0.007)	(0.008)	(0.009)	(0.010)
Average tax bracket			0.046***	0.055***	0.050***	0.049***
			(0.009)	(0.011)	(0.012)	(0.012)
Log employment			0.059***	0.051***	0.057***	0.052***
			(0.005)	(0.006)	(0.003)	(0.007)
State-owned			0.357***		0.209***	0.361***
			(0.013)		(0.013)	(0.021)
Collective-owned			0.093***	0.091***	0.051***	0.071***
			(0.014)	(0.015)	(0.012)	(0.017)
Foreign-invested			0.169***	0.173***	0.170***	0.151***
			(0.030)	(0.031)	(0.030)	(0.034)
HKMT-invested			0.137***	0.142***	0.133***	0.129***
			(0.029)	(0.029)	(0.028)	(0.030)
Profits per worker			-0.002*	-0.001	-0.000	-0.002
			(0.001)	(0.001)	(0.001)	(0.001)
Subsidies per worker			0.008***	0.008***	0.007***	0.007***
			(0.001)	(0.001)	(0.001)	(0.001)
Market share			0.001***	0.002***	0.001***	0.002***
			(0.000)	(0.000)	(0.000)	(0.000)
SOE share			-0.071	-0.061	-0.018	-0.010
			(0.069)	(0.075)	(0.070)	(0.110)
Social security cost			-0.201*	-0.225**	-0.357***	-0.297
			(0.102)	(0.110)	(0.096)	(0.160)
Rural status			-0.070***	-0.064***	-0.045**	-0.040
			(0.020)	(0.022)	(0.018)	(0.025)
Observations	1,022,419	1,022,419	717,570	637,587	590,689	517,243
R-squared	0.005	0.005	0.139	0.105	0.148	0.117

Notes: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors, clustered at the regional level, in parentheses. All estimations are by OLS and include year fixed effects and industry fixed effects at the 2-digit level.

Table 3: OLS estimates of the effects of labor market tightness on firm provision of unemployment insurance. The dependent variable in Columns (1) and (2) is 1 for firms reporting some non-zero insurance payment in a given year, otherwise 0. In Columns (3) and (4), the dependent variable takes the value 1 for firms that joined the UI program in the subsequent year, otherwise 0. Balanced 2001-2005 panel.

	(1)	(3)	(3)	(4)
Labor market tightness	0.005**	0.006***	0.005**	0.003**
	(0.002)	(0.002)	(0.002)	(0.001)
Share of high school		0.119***		
		(0.012)		
Share of college		0.231***		
		(0.020)		
Female share		-0.042**		
		(0.019)		
Union		0.088***		
		(0.008)		
Average wage	0.095***	0.077***		-0.007**
	(0.008)	(0.007)		(0.003)
Average tax bracket	0.044***	0.031***		-0.005**
	(0.006)	(0.006)		(0.002)
Log employment	0.054***	0.052***		-0.005***
	(0.007)	(0.006)		(0.001)
State-owned	0.363***	0.310***		-0.022***
	(0.020)	(0.018)		(0.007)
Collective-owned	0.092***	0.086***		-0.019***
	(0.018)	(0.017)		(0.005)
Foreign-invested	0.183***	0.170***		0.008
	(0.031)	(0.028)		(0.005)
HKMT-invested	0.145***	0.143***		0.016***
	(0.033)	(0.029)		(0.005)
Profits per worker	-0.003***	-0.004***		0.000
	(0.001)	(0.001)		(0.000)
Subsidies per worker	0.005***	0.004***		-0.001
	(0.001)	(0.001)		(0.001)
Market share	0.001***	0.001***		-0.000
	(0.000)	(0.000)		(0.000)
SOE share	-0.006	-0.042		-0.016
	(0.066)	(0.060)		(0.040)
Observations	258,174	258,174	256,381	373,605
R-squared	0.158	0.584	0.172	0.038

Notes: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors, clustered at the regional level, in parenthesis. All estimations are by OLS and include year fixed effects and industry fixed effects at the 2-digit level.

7.2 Additional determinants of UI provision: what can we learn from the control variables?

Interpreting the coefficients of the control variables is not straightforward because the control variables are potentially endogenous in the estimated regressions. Nevertheless, the estimates can be seen as indicative of correlations between a firm's UI provision and its workforce and institutional characteristics, and these could be used to guide future research.

First, it can be noted that the average wage in a firm is positively correlated with the likelihood of UI compliance, and the estimated correlation is quite large. Across specifications, a one percent increase in the average monthly wage, at the average firm, is associated with the firm being 3.4-8.7 percentage points more likely to provide UI. This rather large and positive parameter may seem unintuitive considering that firms may be expected to shift the insurance cost to its workers. In that case, we would expect a negative parameter rather than a positive one. However, the room for such cost-shifting has been shown to be small for Chinese firms (Nielsen and Smyth, 2007). For example, firms with wages close to or at the minimum level do not have the capacity to cost-shift, which could lead such firms to evade payments. This would contribute to a positive correlation between the average wage level and insurance provision.

Looking at the firm size, we see that the parameter estimate has the expected positive sign. One possible reason for this is that it is relatively more difficult for large firms to evade government scrutiny of their activities. As argued by Mares (1997; 2003), the greater market power of large firms should also increase their ability to shift costs over to customers through higher product prices. Some support for this hypothesis is given by the results in Table 2, where we see that the market share parameter estimate is positive and highly significant.

As shown by the descriptive statistics, firm ownership is a highly relevant variable in accounting for variation in UI provision across firms. The estimation results agree with the observation that the highest provision rate is among SOEs, the second highest is for firms with investment from outside the Chinese mainland, and the lowest rates are among collective-owned and private domestic firms. An interesting observation can also be added from an examination of the estimation result using entries, rather than provision, as the dependent variable (Table 3, Column 4). In this case, the estimates show that foreign- and HKMT-invested firms were those most likely to join the system during the period of investigation. Comparing State-owned and private domestic enterprises, we find a more rapid rate of entry in the latter case.

The control variables provide some evidence that a firm's financial position is a determinant of UI provision. The amount of state subsidies received relative to firm size enters with the expected positive parameter, but the ef-

fect is small. We also observe that the total social security cost facing a firm is negatively correlated with its likelihood of provision of UI, and that the effect appears to be quite large.

Lastly, the results can be interpreted as evidence against the exclusion of worker demand as a determinant of UI provision in Chinese firms. The parameter estimate for the tax bracket variable, which increases with the average wage, takes a positive sign, indicating that high wage earners prefer to receive part of their compensation in tax-exempt social insurances rather than in wages. A similar conclusion can arguably be drawn for the positive and highly significant parameter estimates for the education variables. In addition, the parameter estimate for the unionization variable suggests that firms with unions are almost 9 percentage points more likely to provide the insurance.

8. Concluding remarks

Increasing worker coverage of China's unemployment insurance program is of great economic importance not only to Chinese workers. The country's trading partners also stand to gain when a lower household saving rate translates into higher demand for imported goods and a reduction of China's dependence on foreign demand as a source of growth.

The motive of this study was to expand our understanding of the determinants of UI provision in Chinese firms. In particular, the analysis evaluated the effect of labor market tightness on a firm's decision to provide UI. The empirical analysis of firm-level data for 2001-2005 showed that the labor market situation did indeed correlate with firm behavior in the expected way, indicating that tighter labor markets motivated firms to attract potential new employees by providing UI. Although this result is interesting in the sense that it shows that Chinese firms react to labor market conditions, despite the still underdeveloped nature of the country's labor market, the small size of the estimated effect implies we should not over-emphasize this effect. Arguably, future studies should attempt to quantify more precisely this relationship by improving the labor market tightness measure and by considering labor markets of a smaller scope.

Since this study is the first to examine the provision of social insurance in Chinese firms using a highly representative panel dataset, a couple of additional results should be mentioned. Firstly, the estimation results did not support the exclusion of worker demand channels in explaining a firm's decision to provide UI. Variables correlated with demand, such as worker wages, education-level, unionization and firm size, all appeared to explain a sizeable share of the variation in the data. Secondly, the analysis also highlighted the important role of firm ownership, indicating that the political economy of participation could merit more attention in future studies.

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Appendix A

Table A2: Ownership aggregation of the registration-based firm classification

Registration-based classification (since 1998)	Categories for all industrial enterprises from China's Statistical Yearbook
Domestic enterprises	
State-owned	State
Collective owned	Collective
Employee shareholding company	Collective
Joint operation enterprises	
State-owned	State (1)
Collective owned	Mixed
State- and Collective owned	Mixed (2)
Other joint operation enterprises	Mixed (2)
Limited liability companies	
Solely state-owned	State
Other	State or Mixed (3)
Stock companies	State or Mixed (3)
Private enterprises	
Private sole proprietorships	Private
Private partnerships	Private
Private limited liability company	Private
Private stock companies	Private
Other enterprises	Mixed
HKMT-financed enterprises	
Joint equity ventures (JVEs)	HKMT or State (2)
Contractual joint ventures (CJVs.)	HKMT or State (2)
Wholly HKMT-owned	HKMT
HKMT stock companies	HKMT or State (2)
Foreign-financed enterprises	
Chinese-foreign JEVs.	Foreign or State (2)
Chinese-foreign CJVs.	Foreign or State (2)
Wholly foreign-owned	Foreign
Foreign-financed stock company	Foreign or State (2)

Notes: Departures from the CSY aggregation method are reported in footnotes (1)-(3) below. Information about the CSY aggregation methodology is contained in Holz and Lin (2001). 1. Unlike in the Statistical Yearbook, this category is not double-counted as "Mixed" (here termed "Mixed"). 2. When ownership-disaggregated statistics are reported per ownership in the Statistical Yearbook, a proportion of the statistic is double-counted in the "State" category. This proportion corresponds to the aggregate share of the sum of state capital to the sum of total paid-in capital minus individual capital in the registration-based classification category. I instead count an individual firm as State-owned if the state's share in total capital is greater than 50%. 3. This category is counted as "State" in the Statistical Yearbook for firms that are under absolute state control (*guoyou juedui konggu*) or relative state control (*guoyou xiahgdui konggu*). The first implies that the state owns for more than 50% of the total capital. The second that the state holds less than 50% of total capital but that i) its share is relatively large compared to the shares of other ownership categories, or ii) even though one or more other ownership categories have a larger capital share, the state in effect holds the control rights by agreement (*Xiyi kongzhi*). In this paper, only absolute state-controlled firms can be identified, and these are moved from the "Mixed" to the "State" category.

A1. Variable specifications

Average monthly wage: the monthly wage bill of the firm, including bonuses, overtime pay, and social insurance payments deducted from employee wages, divided by total labor.

Average tax bracket: the tax bracket to which the firm's average worker belongs. The sum of one year's deductibles (12*500 RMB) is deducted from the undeflated annual wage per worker, and this net wage is matched to China's progressive personal income tax scale, which has six brackets.

Distance to port: the distance between the firm's prefecture-level city and the nearest port of Hong Kong, Tianjin, or Shanghai.

Distance to market: the distance between the firm's prefecture level city and the nearest big city of Beijing, Taiyuan, Shenyang, Harbin, Nanjing, Wuhan, Guangzhou, Chengdu, Xi'an, Dalian, Shanghai, Chongqing, Changchun, and Shanghai.

Employee skills: the percentage of workers with: i) a high school diploma, and ii) at least a college degree (*Dazhuan* and above) (limited to 2004).

Female proportion: the proportion of females in the workforce (limited to 2004 and 2005).

Industry: dummies for the 2-digit ISIC level.

Market share: Herfindahl index calculated at the 3-digit product level by region.

Profits per worker: firm profits divided by the number of employees.

Rural status: a dummy variable for firms under the administration of a country, small town, street, village, resident-committee or village committee.

Social security cost: Average of 2004 and 2005 levels of the prefecture-city total cost for: i) pension insurance, ii) health insurance, and iii) housing fund payments and housing subsidies.

SOE share: the proportion of workers employed in SOEs by region.

Unionization: a dummy for whether the firm has a union (limited to 2004).

Table A2: Summary statistics of Chinese industrial firms, balanced panel 2001-2005.

	2001	2002	2003	2004	2005
UI participation	0.40	0.41	0.42	0.51	0.54
UI entry (number)		8,455	8,766	14,744	9,824
UI exits (number)		8,038	7,996	7,966	7,472
State-owned (share)	0.24	0.20	0.15	0.11	0.08
Collective-owned (share)	0.24	0.21	0.16	0.09	0.09
Private domestic (share)	0.21	0.27	0.34	0.43	0.46
Foreign-invested (share)	0.07	0.08	0.09	0.10	0.10
HKMT-invested (share)	0.10	0.10	0.10	0.10	0.10
Average Wage (¥)	12,253	12,333	13,104	15,239	16,712
Average tax bracket	1.29	1.34	1.40	1.50	1.60
Employment	394	400	409	407	422
Profit per worker (¥)	14,980	16,930	19,590	27,386	28,564
Rural firms (share)	0.45	0.43	0.41	0.25	0.28
Market share	0.03	0.03	0.03	0.02	0.02
Subsidies per worker (¥)	843	894		1,159	1,251

Notes: Information on firm subsidies is not available for year 2003. "Mixed" ownership not reported.

Table A3: Estimates of the effects of labor market tightness on firm provision of unemployment insurance. The dependent variable is 1 for firms reporting some non-zero insurance payment in a given year, otherwise 0. Unbalanced panel data from 2001-2005.

	Region fixed effects, OLS (1)	Firm fixed effects, OLS (2)	Probit (3)
Labor market tightness	-0.006	-0.003	0.007***
	(0.003)	(0.004)	(0.003)
Average wage	0.084***	0.031***	0.105***
	(0.006)	(0.006)	(0.010)
Average tax bracket	0.042***	0.034**	0.040***
	(0.009)	(0.017)	(0.009)
Log employment	0.062***	0.067***	0.066***
	(0.006)	(0.013)	(0.007)
State-owned	0.345***	0.056***	0.388***
	(0.014)	(0.014)	(0.016)
Collective-owned	0.088***	0.023**	0.106***
	(0.013)	(0.010)	(0.017)
Foreign-invested	0.154***	0.054***	0.180***
	(0.029)	(0.019)	(0.034)
HKMT-invested	0.105***	0.035**	0.146***
	(0.026)	(0.014)	(0.033)
Profits per worker	-0.001	0.003**	-0.002*
	(0.001)	(0.001)	(0.001)
Subsidies per worker	0.008***	0.005***	0.008***
	(0.001)	(0.001)	(0.001)
Market share	0.001***	0.000	0.002***
	(0.000)	(0.000)	(0.000)
SOE share	0.319	-0.563***	-0.092
	(0.210)	(0.159)	(0.079)
Social security cost	-0.018	-0.499	-0.226**
	(0.061)	(0.664)	(0.111)
Rural status	-0.066***	-0.009	-0.075***
	(0.019)	(0.016)	(0.021)
Observations	717,570	717,570	717,570
R-squared	0.153	0.695	

Notes: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors, clustered at the regional level, in parentheses. All estimations include year fixed effects and industry fixed effects at the 2-digit level.

Essay 4: Oil prices and real exchange rate volatility in oil-exporting countries: the role of institutions*

1. Introduction

Empirical research on the role of the oil price as a determinant of the real exchange rates of oil-exporting countries has yielded ambiguous and somewhat puzzling results. While strong relationships between the two variables have been found for some countries, weak or even negative relationships have been found for others.

This paper aims to reconcile this mixed empirical evidence. Drawing on insights from models of the political economy of fiscal spending in countries that produce natural resources, I suggest that co-movements are conditional on a country's legal and political institutions. This argument is clarified in a simple theoretical model, where institutions determine the degree of myopia in state spending of oil revenue. By making spending more balanced over the price cycle, strong institutions cut off the fiscal spending mechanism that causes oil price volatility to spill over to the real exchange rate.

A panel of 33 oil exporters for the period 1985-2005 is used to empirically evaluate the predictions of the model. The key empirical finding is that the tendency of the real exchange rates of these resource-exporting states to co-move with the oil price is conditional on their institutions. In particular, high bureaucratic quality and strong and impartial legal systems are found to lead to currencies that are less affected by oil price fluctuations. These results offer an explanation to the ambiguous findings in the empirical literature on real exchange rate determination in oil-exporting countries. They also indicate that a sound institutional setup can prevent the negative effects on growth from a volatile real exchange rate (Bagella et al., 2006; Aghion et al, 2009).

The paper is organized as follows. Section 2 presents previous research on oil prices and real exchange rates. Section 3 lays out the theoretical framework and section 4 provides information on the data and the economet-

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ric specification. Empirical results are contained in section 5, and section 6 concludes.

2. Evidence on oil prices and real exchange rates

In studies of real exchange rate determination in oil-exporting countries terms of trade are commonly approximated by the real oil price (Baxter and Kourparitsas, 2000; Backus and Crucini, 1998), and some have used labels such as "petrocurrency" to describe the perceived importance of this factor in explaining real exchange rate movements. Empirical evidence has however been inconsistent. While changes in the oil price appear to trigger currency movements in some countries, there seems to be little evidence for that relationship for some of the biggest oil exporters in the world.

Among the studies that document an important role for the oil price real exchange rate determination is Korhonen and Juurikkala's (2007) study a panel of nine OPEC countries[1]. In country-specific contexts, Zalduendo (2006) Koranchelian (2005) and Mongardini (1998) document a key role of the oil price as a trigger of real exchange rate movements in Venezuela, Algeria and Egypt respectively. Several studies also provide empirical evidence in favour of the Russian Rouble being an petrocurrency (Spatafora and Stavrev, 2003; Oomes and Kalacheva, 2007).

Contrasting these findings, researchers have reported statistically insignificant, or numerically weak, relationships between the Norwegian Krone and the oil price (Bjørvik et al., 1998; Bjørnland and Hungnes, 2008; Akram 2000; 2004). Similarly, there is little evidence that the Canadian dollar is a petrocurrency, with researchers again reporting insignificant (Gauthier and Tessier, 2002) or even negative relationships (Amano and van Norden, 1995). Finally, in a study of the world's largest oil exporters, Russia, Norway and Saudi Arabia, Habib and Kalamova (2007) find that the oil price influences the movements of the Russian rouble, but that the currencies of major oil producers Norway and Saudi Arabia remain unaffected by price volatility.

In single-country settings, some attempts have been made to find explanations for variations over time in the degree of covariation between real exchange rates and the real oil price. Sosunov and Zamulin (2006) and Issa et al. (2006) point to the relative importance of oil exports in the domestic economy to account for the degree of appreciation following oil price hikes. Habib and Kalamova (2007) informally discuss the potential importance of policy responses and revenue management.

[1] Algeria, Ecuador, Gabon, Indonesia, Iran, Kuwait, Nigeria, Saudi Arabia and Venezuela.

3. Theoretical framework

This section contains a stylized model of real exchange rate determination in an oil-exporting economy. The model is deliberately kept simple and its sole purpose is to motivate the empirical specification that will be used to test the main hypothesis, namely that institutions mediate the effect of the oil price on the real exchange rate. Consider a small open economy producing oil, a tradable and a non-tradable good. The non-tradable good is consumed domestically by the state and by private consumers, while the tradable good is consumed by domestic and foreign consumers. The oil producing sector is owned by the government and employs a negligible share of the domestic labor force. Oil and tradables are sold on the world market at exogenously given prices.

3.1 Production

The production of non-tradable ($Y_{N,t}$) and tradable goods ($Y_{T,t}$) at time t is given by $Y_{i,t} = A_{i,t} L_{i,t}^{\alpha}$ where $0 < \alpha < 1$, A_i is a productivity factor, and L_i is the labor input in sector $i = N, T$. Labor is fully flexible, which implies that wages equalize between the two sectors and that the exogenous labour supply equals the sum of labor demand $L_t = L_{N,t} + L_{T,t}$. Profit maximization yields the first order conditions $w_t = P_{i,t} A_{i,t} \alpha L_{i,t}^{\alpha-1}$ which can be combined to give an expression for the relative supply of non-tradable and tradable goods as

$$\frac{Y_{N,t}}{Y_{T,t}} = \left(\frac{A_{N,t}}{A_{T,t}}\right)^{\frac{1}{1-\alpha}} \left(\frac{P_{N,t}}{P_{T,t}}\right)^{\frac{\alpha}{1-\alpha}}. \tag{1}$$

3.2 State and private consumption

The state obtains income from taxing the tradable sector and from selling oil (M_t) on the global market at the price $P_{O,t}$ [2]. In each time period, the state's

[2] The IMF (2007) reports the ratios of oil revenues to total fiscal revenues as five year-averages for the 2000-20005 period for several of the countries included in this study, namely Algeria (70.5), Angola (79.8), Azerbaijan (33.3), Cameroon (27.7), Colombia (10.0), Congo (Brazzaville) (69.9), Equatorial Guinea (85.2), Gabon (60.1), Indonesia (30.3), Iran (65.5), Kazakhstan (25.1), Kuwait (74.7), Libya (80.2), Mexico (33.3), Nigeria (78.9), Norway (24.0), Oman (83.4), Qatar (68.4), Russia (19.5), Saudi Arabia (83.1), Sudan (49.8), Trinidad and Tobago (36.4), UAE (66.1), Bolivia (20.9).

consumption of non-tradables G_t is the sum of income received from taxes and a fraction λ of total oil revenue. Assuming that the income gained from taxes corresponds to a fraction τ of the income of the tradable sector $P_{T,t}Y_{T,t}$, the government's consumption of non-tradables can be expressed as[3]

$$P_{N,t}G_t = \tau P_{T,t}Y_{T,t} + \lambda P_{O,t}M_t \tag{2}$$

where $0 < \lambda < 1$. Since oil revenue will not last forever, we interpret the parameter λ as the degree of myopia in revenue spending. A government with a high λ will consume a large share of a sudden oil revenue increase, while a government with low λ will save a larger share of that income, letting a smaller share spill over on the consumption of non-tradables. The share of total oil revenue not spent on non-tradables $(1-\lambda)P_{O,t}M_t$ is invested by the government at the international credit market at the fixed interest rate r. The government's budget constraint is hence such that investment in period $t+1$ will equal income from investment[4], oil revenue and taxes in period *t*, minus the amount spent on non-tradables in the same period.

Turning next to the consumers in the economy, these actors own the firms, get income in the form of wages and profits, pay taxes and derive utility from consumption of non-tradable and tradable goods. To simplify the model, it is assumed that consumers are restricted from borrowing on the credit market. The budget constraint of consumers is

$$P_{N,t}Y_{N,t} + (1-\tau)P_{T,t}Y_{T,t} = P_{N,t}C_{N,t} + P_{T,t}C_{T,t}\,. \tag{3}$$

Assuming Cobb-Douglas utility with weights γ and $1-\gamma$, a fraction γ of total consumer expenditure is spent on non-tradables, i.e.

$$P_{N,t}C_{N,t} = \gamma(P_{N,t}C_{N,t} + P_{T,t}C_{T,t})\,. \tag{4}$$

We can now derive an expression for the relative *demand* for tradables and non-tradables. This is achieved by combining the budget constraints for the state, and for consumers, substituting for $C_{N,t}$ using the equilibrium condition for the market for non-tradable goods[5] and by expressing all terms in the equation as shares of the size of the economy for non-tradable goods by dividing them with $P_{T,t}Y_{T,t}$. After some algebra, we get that

[3] The assumption that only the tradable sector is taxed is made to simplify the derivations.
[4] Where investment in the first period is set to zero.
[5] $Y_{N,t} = C_{N,t} + G_t$

$$\frac{Y_{N,t}}{Y_{T,t}} = \frac{1}{1-\gamma}\left[\gamma(1-\tau)+\tau+\lambda\left(\frac{P_{O,t}M_t}{P_{T,t}Y_{T,t}}\right)\right]\frac{P_{T,t}}{P_{N,t}} \tag{5}$$

3.3 The Foreign Economy

Production technology, private consumption and the labor market in the foreign economy are identical to that in the domestic economy, but the foreign economy lacks an oil sector. Using * to denote the foreign economy, we can derive analogous equations (1')-(5').

3.4 Real Exchange rate Determination

The real exchange rate of the oil-exporting economy (Q_t) is defined as the price in the foreign currency of a domestic basket of consumption relative to the price of a foreign basket of consumption $Q_t = E_t P / P_t^*$, where * denotes the foreign economy and E_t is the price of domestic currency. An increase in Q_t hence implies real appreciation. We assume that the domestic and foreign price levels are geometric averages of the prices of traded and non-traded goods with weights $1-\gamma$ and γ respectively. We can then write the aggregate price level such that $P_t = P_{N,t}^{\gamma} P_{T,t}^{1-\gamma} = P_{T,t}\left(P_{N,t}/P_{T,t}\right)^{\gamma}$ and the real exchange rate expression can be rewritten as

$$Q_t = \frac{E_t P_t}{P_t^*} = \frac{E_t P_{T,t}}{P_{T,t}^*}\left(\frac{P_{N,t}/P_{T,t}}{P_{N,t}^*/P_{T,t}^*}\right)^{\gamma} \tag{6}$$

The law of one price is assumed to hold for the tradable good so that $P_{T,t}^* = E_t P_{T,t}$. Next, equations (1) and (5), describing the relative demand and supply of tradable and non-tradable goods, are combined for both the domestic and the foreign economies. This enables us to derive the price ratios for tradable and non-tradable goods in the two economies, and inserting these ratios into (6) we get the following equation for the real exchange rate of the oil-producing country:

$$Q_t = \left[\frac{A_{T,t}/A_{N,t}}{A^*_{T,t}/A^*_{N,t}}\right]^{\gamma} \left[\frac{\gamma(1-\tau)+\tau+\lambda\dfrac{P_{O,t}M_t}{P_{T,t}Y_{T,t}}}{\gamma^*(1-\tau^*)+\tau^*}\right]^{\gamma(1-\alpha)} \tag{7}$$

The first factor $\left[\frac{A_{T,t}/A_{N,t}}{A^*_{T,t}/A^*_{N,t}}\right]^{\gamma}$ is the standard Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964) whereby increased productivity in the domestic tradable goods sector leads to real appreciation of the domestic currency. From this equation we get two important insights which will guide our empirical specification. The first is that what should matter for the real exchange rate is *oil revenue relative to the size of the economy*. An increase in oil revenue should have a similar effect independent of whether it arises as a result of a price increase or as a result of increased production. Second, the effect of an increase in oil revenue is contingent on the degree of myopia of the government. The more myopic the government is, the more it spends of an increase in revenue, and the stronger is the spill-over of the oil price movement over on the real exchange rate. Moreover, the bigger the country's oil revenue is, as compared to the size of the size of the economy at large, the more important is the effect of myopia. Hence, there should be an *interaction effect* between income and variables meant to reflect the degree of myopia in the government's spending behaviour. Because myopic governments spend a large share of the incoming revenues, the real exchange rates of these countries will co-move with the oil income. In other words, volatility of the oil price will be reflected in volatility of the real exchange rate. In this paper, we think of the myopia parameter λ as reflecting the institutional setting in which the policy maker operates and which affect his or her spending incentives.

4. Empirical specification and variable selection

The next step is to construct an empirical specification of the real exchange rate equation (7). The key parameter of interest λ, is assumed to depend on a vector of governance characteristics $X_{j,i,t}$ according to

$$\lambda_{i,t} = \sum_{j=0}^{K} \beta_j X_{j,i,t}$$

where the sub-script *i* denotes the country, and *j* = 1, 2,...,K indicates legal and political institutions that influence the spending behaviour of the policy

maker. The first element in the vector $X_{j,i,t}$ is set to one to test the unconditional effect of oil dependency on the real exchange rate (see Issa et al. 2006; Sosunov and Zamulin, 2006). Taking logs of Equation (7) gives the test equation

$$\ln Q_{i,t} = \alpha_0 + \alpha_2 \ln PR_{i,t} + \sum_{j=2}^{K} \beta_j S_{i,t} X_{i,t,j} + \varepsilon_{i,t} \tag{8}$$

where, $Q_{i,t}$ is the real exchange rate in country i at time t, PR denotes the productivity differential between traded and non-traded goods relative to the rest of the world, S represents the relative size of the oil sector in the domestic economy and a random disturbance term ε has been added.

4.1 Institutional variables

Government consumption impacts on the real exchange rate through its bias toward nontradables over tradables (De Gregorio et al., 1994; Ricci et al. 2008). Case studies of the fiscal policies of natural resource rich countries, and oil-exporters in particular, have pointed out myopic spending of revenue as the major policy flaw (Gelb, 1988; Auty, 2001). Another insight from the literature is that a country's institutional setup is an important determinant of its spending decisions.

In the field of political economy, the theoretical literature has proposed two main mechanisms through which institutions affect the degree of responsibility in State spending of natural resource windfalls. One strand of the literature focuses on rent seeking as the mechanism, while the other strand focuses on patronage. These two strands also lead to different conclusions about what types of institutions it is that matter.

Let us first examine the model that focuses on the role of institutions in determining the incentives for "rent-seeking" among powerful social groups attached to state spending. Lane and Tornell (1999) argue that a "strong legal-political institutional structure" prevents resource rents from being systematically re-directed to the informal sector. Mehlum et al. (2006) offer a similar model on rent-seeking, but focus their attention on the incentives for private entrepreneurs. Essentially, they show that if the institutional framework promotes the profitability of private enterprise, entrepreneurs choose to engage in production rather than rent seeking. In turn, the government's spending pattern is smoothed out over the revenue cycle. Empirically, Mehlum et al. (2006) operationalize the institutional environment thought to prevent rent-seeking by an index measuring law and order, bureaucratic quality, and contract enforcement (using the same specification as Knack and Keefer, 1995).

Next, we have the model that focuses on political patronage (Robinson et al., 2006). It builds on the intuition that as governments seek re-election, one way of achieving this goal is to shore up support via the provision of political appointments and funds. In an institutional setting that promotes accountability, diverting resource rents for the purpose of patronage cannot be done. However, in the absence of such institutions, resource revenues can be systematically diverted towards patronage. Empirical applications of the patronage model have used the democracy index from the World Bank's Polity IV dataset to measure political accountability (e.g. Kolstad, 2009).

Drawing on systematic case studies of oil-exporting economies with varying institutional settings, Eifert et al. (2002) seek to identify institutional characteristics that have promoted anti-cyclical spending of windfalls. Using argument of both rent-seeking and patronage, they propose four characteristics.

Firstly, the degree of stability of the political framework of the country's party system is expected to influence spending incentives. In stable systems, political power is derived via economic performance. Unstable systems on the other hand, feature competition for power, often via interventions from military or other interest groups. This gives the ruling elites incentives for short-term spending to gain these groups' support rather than holding a long-run policy horizon of growing the economy. A competent bureaucracy is also a key factor for a stable political framework. In their presence, policy tends to be stable despite shifts in the government in power, and special interests cannot influence spending via the bureaucracy.

The second characteristic is the degree of social consensus. The importance of this institution lies with the role of the public as a counter-weight to irresponsible spending behaviour of the State. In countries with a high degree of social consensus, the public supports long-term spending horizons and transparency in the budget process.

Third, Eifert et al. (2002) highlight the importance of the way governments (or aspiring governments) obtain or maintain support. If the power of the ruling elites is gained via patronage, oil rents are part of the political game rather than a way to achieve public support via long-term investments.

The fourth institutional characteristic is the role of state institutions in underpinning markets and the distribution of rents. Here, clear property rights and the impartiality of the judicial system are key factors. A de-personalized functioning of markets determines the relative pay-offs for interest groups to engage in rent-seeking groups or productive activities.

Institutional data that can capture the institutional characteristics outlined in the literature is obtained from the Political Risk Services, a private company that assesses economic and political risk across a large number of

countries.[6] In total, eight variables are selected. To reflect the institutions expected to reduce the incentives for rent-seeking, I follow the literature and include a measure for the quality of the legal system, and one for the quality of the bureaucracy (e.g. Kolstad, 2009). A measure of democratic accountability is included to reflect restraints on patronage. Variables for government stability, corruption, and the occurrence of military interventions are used to capture the stability of the party system and the political framework, as argued for by Eifert et al. (2002). The country's investment profile, measuring the degree of contract viability, profits repatriation and payment delays, is added as a proxy for the market environment. Finally, an attempt is made to capture the degree of social consensus by including an index of "socioeconomic conditions", which includes information on unemployment, consumer confidence, and poverty. More thorough descriptions of the institutional variables are provided in Table A2.

Figures A1 and A2 show the distributions of the ordinal institutional variables. Regression analysis using a dataset that includes ordinal covariates with relatively few categories calls for caution when choosing an econometric technique. Erroneously treating ordinal variables as continuous may yield misleading results, unless the effect of the ordinal variable is linearly related to its categorization (Jöreskog, 1994).[7] If this assumption is accepted, the variable can regarded as being measured on an interval scale where the step between each category on the ordinal variable is equally large. A test is constructed by decomposing the ordinal variables into dummy variables for each category value. For each set of dummy variables $i = 1, 2, ..., N$, the restriction $\beta_1 = (1/2)\beta_2 = (1/3)\beta_3 = (1/N)\beta_N$ is thereafter tested in an estimated fixed effects model.[8] These tests show that the indicators for Bureaucratic Quality, Corruption, Democratic Accountability, Law and Order, and Investment Profile may be assumed to be ordinal when interacted with the ratio of oil exports to GDP. In contrast, this is not a reasonable assumption for the variables measuring Government Stability, Military involvement in Politics and Socioeconomic Conditions.[9] The latter variables are hence not treated as measured on an interval scale, but instead transformed into two

[6] These data were introduced as institutional measures in economic research by Knack and Keefer (1995) and are now widely used (e.g. Hall and Jones, 1999; Acemoglu et al., 2001).

[7] In other words, that a move between categories 0 to 1 on an ordinal variable measuring, for example, bureaucratic quality, leads to the same effect on the dependent variable as a move between categories 5 to 6.

[8] In the ICRG data, countries are sometimes given half points, placing them between two categories. Before restrictions are tested, these observations are sorted into the closest category. Sorting them into the higher or lower category does not affect the conclusions of the restriction tests.

[9] P-values of the Wald tests are as follows, Bureaucratic Quality: 0.25; Corruption: 0.16; Democratic accountability 0.21; Investment Profile: 0.48; Law and Order: 0.24; Military in Politics: 0.02; Socioeconomic Conditions: 0.00; Government Stability: 0.08.

threshold variables each, one for the top quartile of the distribution, and one for the bottom quartile.

4.2 Continuous time series variables

The remaining variables in equation (8) are constructed as follows. For the real exchange rate Q, the IMF:s CPI deflated real effective exchange rate index is available for 17 of the countries in the sample. For the remaining 16, I calculate real bilateral exchange rate indices vis-à-vis the US[10]. A higher value of the real exchange rate index implies real appreciation. Next, the real average oil price series P_O is derived by deflating the average price of crude oil in US dollars by the US consumer price index with base year 2000. The real exchange rate measures and the real oil price series are transformed to natural logarithms.

As a proxy for the size of the oil sector in the domestic economy, denoted S in equation (8), I first calculate the share of oil exports to GDP by first multiplying volume of net oil exports by the oil price in current USD for each year. This measure is then divided by total GDP for the exporting country.

To control for the Balassa-Samuelsson effect, corresponding to PR in the test equation, I include the log-difference between the oil exporters' per capita GDP in PPP-based constant 2000 US dollars and per capita GDP in the US. The use of this proxy rests on the assumption that the productivity advantage of high income countries is primarily found in the tradable rather than the non-tradable sector.[11] Additional variable information is placed in Table A2.

The time period 1985-2005 is selected in order to maximize data availability. For this period, data from the EIA on net oil exports is used to identify 38 countries with positive net export values for at least half of the years. Of these countries, Iraq, Turkmenistan and the United Arab Emirates are excluded due to the unavailability of consistent time series for the real exchange rate or GDP. Furthermore, Brunei and Equatorial Guinea are excluded due to the unavailability of some institutional and political data. In the resulting sample of 33 countries[12], net oil imports are recorded in less

[10] In the sub-sample of 17 countries for which both measures are available, the correlation coefficient between the real effective exchange rate index and the bilateral exchange rate index is 0.78.

[11] Using aggregated production to approximate for productivity differences in traded and non-traded goods in the domestic and foreign economies is standard practise when facing the common problem of data unavailability for the more exact measures. Related research resorting to this method includes studies by Koranchelian (2005) and Korhonen and Juurikkala (2007).

[12] Algeria, Angola, Argentina, Azerbaijan, Bolivia, Cameroon, Canada, Colombia, Congo (Kinshasa), Congo (Brazzaville), Egypt, Gabon, Indonesia, Iran, Kazakhstan, Kuwait, Libya,

than 3 percent of all years. The final panel consists of 551 observations (see Appendix, Table A1 for a record of missing data)

Concerning the choice of econometric model, macroeconomic time series may contain unit-roots, in which case the use of OLS can produce invalid estimates. To choose the appropriate econometric technique, it is therefore important to establish if the macroeconomic time series used, namely the real exchange rate measure ($Q_{i,t}$), the real oil price ($P_{O,t}$), the productivity differential ($PR_{i,t}$) and the ratio of oil exports to GDP ($S_{i,t}$)[13] are stationary or not. Four panel unit root tests are carried out for this purpose. The Levin, Lin & Chu (2002) test is employed to test for a common unit root process in the panel as a whole, while the Im, Pesaran & Shin, the ADF- and the PP-tests allow the AR coefficients to differ between the sampled countries. Table 1 summarizes the results from four panel unit root tests, which show coherent rejection of the null-hypothesis of non-stationarity for all three series and both test types.

The real oil price time series is examined using four different unit root tests (data plot in Appendix, Figure A3). Results indicate a stationary series. The DFGLS and PP tests reject the null of unit root at one and ten percent respectively, while the KPSS test fails to reject the null of stationarity at the one percent level. Thus, according to the unit root test results, the four macroeconomic variables used in the analysis are found to be stationary. This allows estimation using a panel data regression model with country fixed effects, a model which efficiently controls for all unobservable and time-invariant country characteristics.

Malaysia, Mexico, Nigeria, Norway, Oman, Papua New Guinea, Qatar, Russia, Saudi Arabia, Syria, Trinidad and Tobago, Tunisia, United Kingdom, Venezuela, Vietnam and Yemen

[13] Since exports constitute a share of GDP, the two series are cointegrated by default. The ratio of exports to GDP variable used in this paper, which is restricted to taking on values within the interval 0-1, is hence stationary by definition. This is confirmed by the unit root test results.

Table 1: Panel unit root tests.

	Im, Pesaran and Shin	Augmented Dickey Fuller	Phillips-Perron
Real Exchange Rate (Q)	-2.04 (0.02)	82.69 (0.08)	211.71 (0.00)
Productivity Differential (PR)	-1.04 (0.15)	89.04 (0.02)	78.19 (0.11)
Exports to GDP Ratio (S)	5.73 (0.00)	174.65 (0.00)	253.07 (0.00)

Notes: The table reports results of panel unit root tests, all of which test the null-hypothesis of unit root. P-values are reported in parenthesis. The test statistics correspond to the w-stat in Im, Pesaran and Shin's (2003) test together with the Fisher Chi-square statistic in the ADF- and PP-tests for individual unit root processes. Lag-lengths are selected according to the Schwartz criterion and all tests include a constant but no trend.

5. Estimation results

Table 2 displays estimates of Equation (8) using a panel data model with country fixed effects.[14] Looking first at the top row, we see that the coefficient on the productivity differential is positive and highly significant. A higher level of GDP vis-à-vis that of the foreign economy triggers appreciation, which is consistent with the Balassa-Samuelson effect. In row two we find the parameter for oil dependency variable. It is positive and statistically significant at the one percent level, indicating that higher oil exports relative to total GDP yields appreciation of the real exchange rate.

The remaining rows of the table contain the parameter estimates for the institutional variables interacted with the resource wealth measure. All institutional indicators are standardized to have a mean of zero and a standard deviation of one. Five of the eight variables are statistically significant and enter with the expected signs. By comparing the coefficient sizes, we can determine their order of importance for the oil price real exchange rate relationship. The largest coefficient corresponds to the indicator of bureaucratic quality. To interpret the negative sign of this parameter, we recall that all institutions are measured so that higher values indicate stronger legal and political settings, and that the indicators are normalized with mean zero. This means that the coefficient on oil exports captures the effect of oil revenue on

[14] When data is collected at country level, spatial dependence of observations can arise. This is because the observed units are not randomly selected from a large population, which increases the likelihood of correlation between outcomes from adjacent units. To address this concern, covariance matrixes are estimated using the spatial correlation consistent covariance matrix estimator devised by Driscoll and Kraay (1998). This estimator is also robust for heteroskedasticity and autocorrelation, the other two main sources of inconsistent estimates in panel data estimation.

the real exchange rate for a country with mean bureaucratic quality. The negative parameter for the interaction term between oil exports and bureaucratic quality shows that the effect of oil revenue on the real exchange rate is *lower* for countries with relatively good bureaucratic quality. This result is in line with the theory that bureaucratic quality makes government behaviour less myopic, so a lower fraction of the increase in oil revenue is spent and the real appreciation following a price increase is smaller.

The second most important institutional variable is the index of law and order. Analogous to bureaucratic quality, the negative sign on this interaction shows that countries with strong legal systems can better insulate their currencies from volatility in oil revenues. We can interpret this relationship as follows. High-quality legal systems can create and enforce clear guidelines for the distribution of resource rents. They also increase the profitability of entrepreneurial activities relative to rent-seeking. As a result, pressure for short-term spending of revenue can be expected from special interest groups in the oil-exporting country.

The estimates also show a negative and significant parameter for the interaction with the indicator of high government stability. This effect is in accordance with our expectation that stability of the political system supports long policy horizons and responsible spending of oil revenue.

Next, we focus on the role of military involvement in the political system, and note a negative and significant parameter estimate for countries with high military involvement. Possibly, high military involvement is associated with more government stability and a reduction of business risk.

The last significant institutional variable is the one measuring socioeconomic conditions. In the total sample, the results show that bad conditions, that is, a high level of poverty and unemployment, is associated with more real exchange rate appreciation following an oil price upturn. This result is in accordance with Eifert et al.'s (2002) conjecture that a lower degree of social consensus reduces the level of constraint imposed on state spending from the public, and thus increases the likelihood of irresponsible, i.e. pro-cyclical, spending.

Finally, the estimation results suggest that the effect of the oil price on the currencies of oil-exporters is not conditional on corruption, democratic accountability or investment profile.

Table 2: Effects of the oil price on real exchange rates, conditional on legal and political institutions.

	(1)	(2)	Predicted Sign
Productivity Differential	0.687*** (.138)	0.718*** (.133)	+
Ratio of Oil Exports to GDP	0.110* (.060)	0.124*** (.045)	+
Institutional variables interacted with the ratio of oil exports to GDP			
Bureaucratic Quality	-0.172*** (.048)	-0.162*** (.045)	-
Corruption	0.033 (.075)		-
Democratic Accountability	-0.034 (.043)		-
Investment Profile	-0.018 (.013)		-
Law and Order	-0.160*** (.040)	-0.181*** (.039)	-
Low Government Stability	-0.027 (.021)		+
High Government Stability	-0.065*** (.019)	-0.061*** (.023)	-
High Military Presence in Politics	-0.043* (.015)	-0.042*** (.019)	+
Low Military Presence in Politics	-0.011 (.016)		-
Bad Socioeconomic Conditions	0.075*** (.028)	0.072*** (.027)	+
Good Socioeconomic Conditions	0.025 (.017)		-
R-squared	0.20	0.20	

Notes: The dependent variable is the combined time series of the real effective exchange rate and bilateral real exchange rate index variables in logs. Institutional variables are measured so that higher values indicate stronger/better institutions, and are standardized with mean zero and standard deviation one. The estimation includes country-specific fixed effects, and Driscoll-Kraay standard errors are reported in parenthesis: *, significant at 10%; **, significant at 5%, *** significant at 1%.

In Column 2, Table 2, we exclude the non-significant variables from Column 1. The estimation results remain largely unaltered. Next, we include the un-interacted oil price and the institutional variables. Results are places in Table A4. We see that the parameter on the oil price is small and not statistically significant. With a few exceptions, the un-interacted institutional variables are also statistically irrelevant, while all the interaction terms from Column 2, Table 2, sustain their signs and significance in the expanded specification. The exception to this observation is the indicator for democratic accountability. After controlling for the fact that countries with high values on this variable have more appreciated real exchange rates (as indicated by the positive and significant parameter on the un-interacted indicator), the parameter on the interaction term takes the expected negative sign and is highly significant. This result indicates some support for the patronage model, which emphasizes the role of accountability for reducing irresponsible revenue spending to buy political support. Further robustness tests are carried out by excluding Western countries, Middle Eastern countries and African countries respectively. Table A4 in Appendix shows that the results are reasonably robust to these changes, in particular the effects associated with bureaucratic quality and the legal system.

We can now use the estimation results to draw some conclusions about the performance of individual oil-exporters. Figures 1 and 2 show the average country scores for the indicators of bureaucratic quality and law and order. Importantly, although perhaps not surprisingly, the western mature democracies Norway, Canada, and the UK have the highest average scores on both indicators, while predatory and non-democratic African regimes record the lowest. In the case of law and order, the high scores of Middle Eastern states are also noteworthy. In particular, Saudi Arabia places fourth after the western democracies.

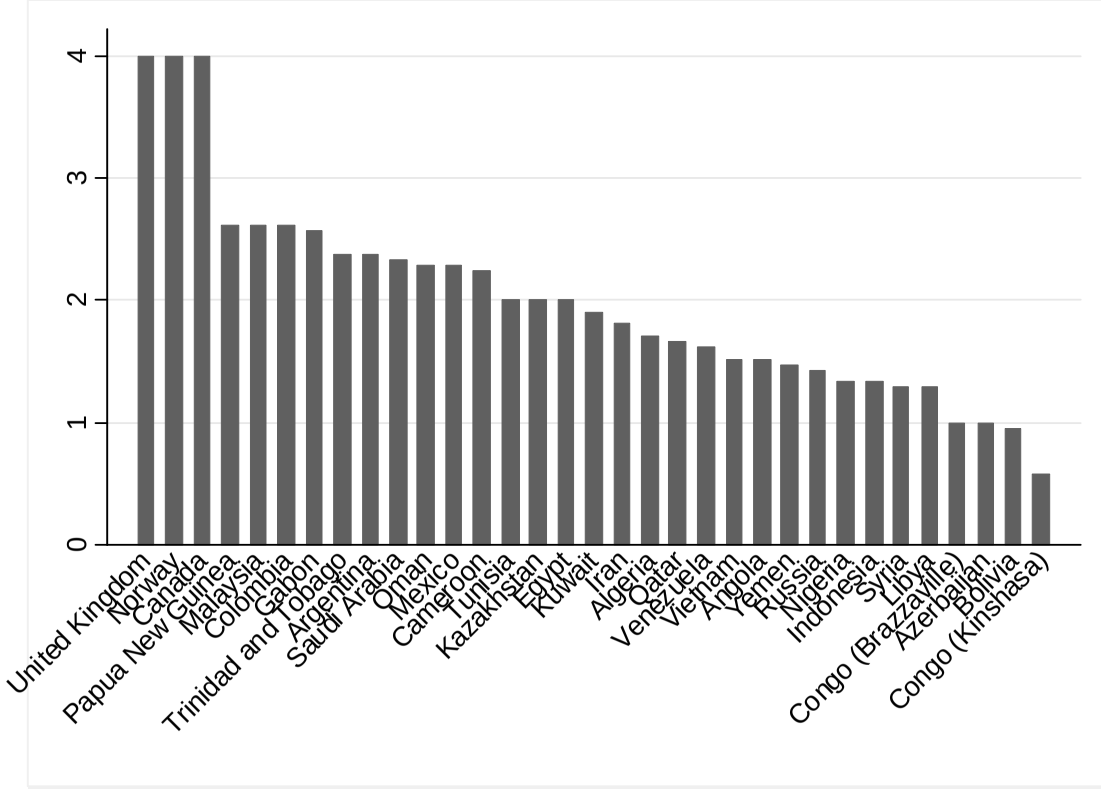

Figure 1: Average Bureaucratic Quality score by country, 1985-2005.

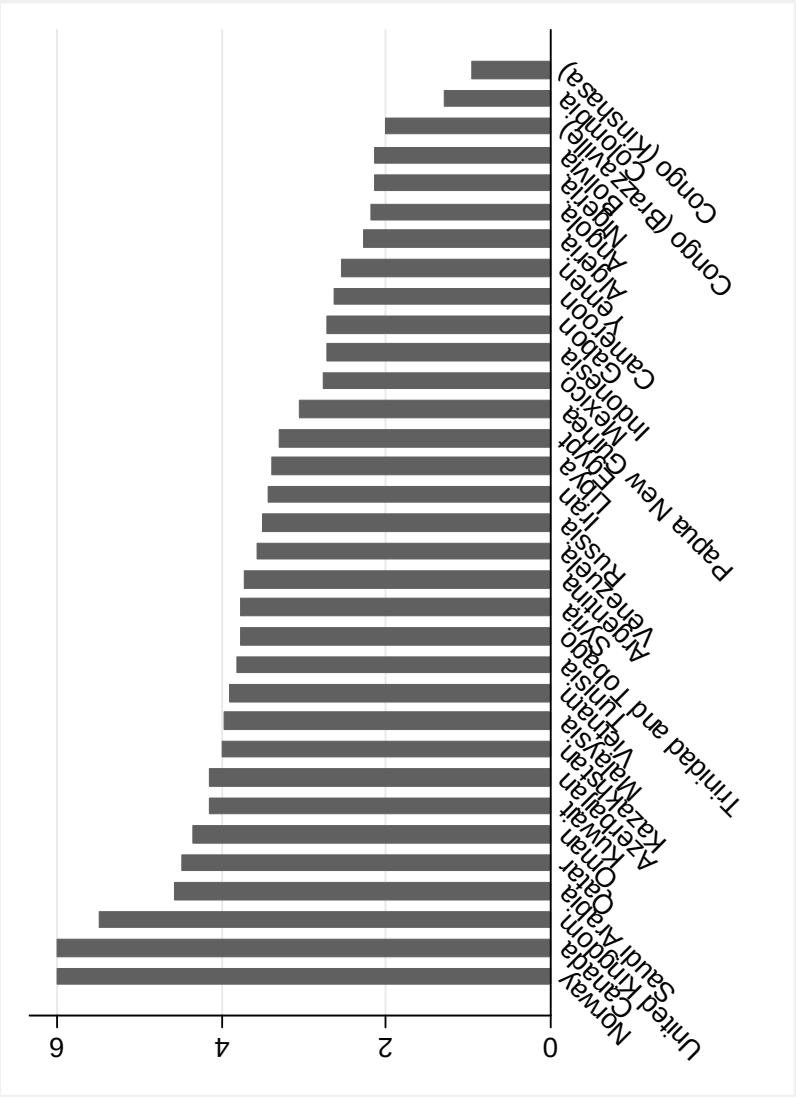

Figure 2: Average Law and Order score by country, 1985-2005.

6. Conclusions

This paper highlighted the role of good political and legal institutions for the relationship between the oil revenue and the real exchange rates of oil-exporting countries. A simple theoretical model was used to show that the effect of oil price movements on the real exchange rate of an oil-exporting economy depended on the degree of myopia in government spending of oil revenue. The empirical relevance of eight governance indicators, believed to affect the spending behaviour of governments, was evaluated on a panel of 33 oil exporters for the period 1985-2005.

The main empirical finding is that the co-variation between the oil price and the real exchange rates of the sampled oil exporters is conditional on political and legal institutions. In particular, currencies in countries with strong bureaucracies and legal systems are less affected by change in oil revenue. This finding is in line with political economy models of rent-seeking (Mehlum et al. 2006; Tornell et al., 1999). However, it adds to the literature by highlighting the role of the real exchange rate as a mediator of the economic effect of institutions in resource dependent countries.

The results in the paper suggest that oil-exporting countries with sufficiently strong institutions can avoid the resource curse associated with a volatile real exchange rate. In doing so, they also provide an explanation for the ambiguous evidence in the empirical literature on real exchange rate determination in oil-producing economies. The lack of strong co-movements between the oil price and the real exchange rate, even in the cases of heavily oil-dependent economies such as Norway, Canada or Saudi Arabia, may be the consequence of favourable institutional characteristics in these countries.

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Appendix A

Table A1: Missing data for continuous variables.

	Real Effective Exchange Rate	Real Bilateral Exchange Rate	Productivity Differential	Ratio of Oil Exports to GDP
Congo (Kinshasa)				85-94
Algeria				97-2001
Iran				91,92
Argentina	X			
Egypt	X			
Indonesia	X			
Mexico	X			
Congo (Brazzaville)	X	97		
Angola	X	85-89		
Syria	X	03,04,05		
Kuwait	X		90,…,94	
Oman	X	85-99	2005	2005
Vietnam	X	85-94,04		85-95
Qatar	X		X	84-93,00,01
Libya	X	2005	X	88,89
Yemen	X	-89	-89	-89

Note: Intervals indicate years for which data is missing, while an "X" indicates unavailability for all years.

Table A2: Variable descriptions and sources.

Real Effective Exchange Rate Index (Q): CPI deflated real effective exchange rate index, base year 2000 = 100 (in logs). Source: IFS, IMF.

Real Bilateral Exchange Rate Index (Q): Real bilateral exchange rate index vis-à-vis the United states, base year 2000 = 100 (in logs). Source: WDI, WB.

Real Oil Price (P): Average oil price in current USD deflated by US CPI, base year 2000 = 100 (in logs). Source: IFS, IMF.

Productivity Differential (PR): Log-difference between GDP per capita in PPP-based constant 2000 USD in the country and per capita GDP in the US. Source: WDI, WB.

Ratio of Oil Exports to GDP (S): Ratio of net oil exports (Source: EIA), multiplied by the average oil price (Source: IFS, IMF), to total GDP (Source: WDI, WB).

From Political Risk Services

Bureaucratic Quality (0-4): Bureaucratic strength and expertise to govern without drastic changes in policy or interruptions in government services. Autonomy of the bureaucracy from political pressure and the existence of mechanisms for recruitment and training of bureaucrats.

Corruption (0-6): Corruption within the political system, specifically excessive patronage, nepotism, job reservations, 'favour-for-favours', secret party funding, and suspiciously close ties between politics and business.

Democratic Accountability (0-6): Responsiveness of the government to its people, where points are awarded based on types of governance rated from the lowest number of points to the highest as i) alternating democracy, ii) dominated democracy, iii) de-facto one-party state, iv) de jure one-party state, and v) autarchy.

Government Stability (0-12): Sum of: government unity (0-4); legislative strength (0-4); and popular support (0-4).

Investment Profile (0-12): Sum of: contract viability/expropriation (0-4); profits repatriation (0-4); and payment delays (0-4).

Law and Order (0-6): Sum of: strength and impartiality of the legal system (0-3) and popular observance of the law (0-3).

Military in Politics (0-6): Military involvement in politics, grading instances of involvement ranging from those caused by internal or external threat to a full-scale military regime.

Socioeconomic Conditions (0-12): Sum of: unemployment (0-4); consumer confidence (0-4); and poverty (0-4).

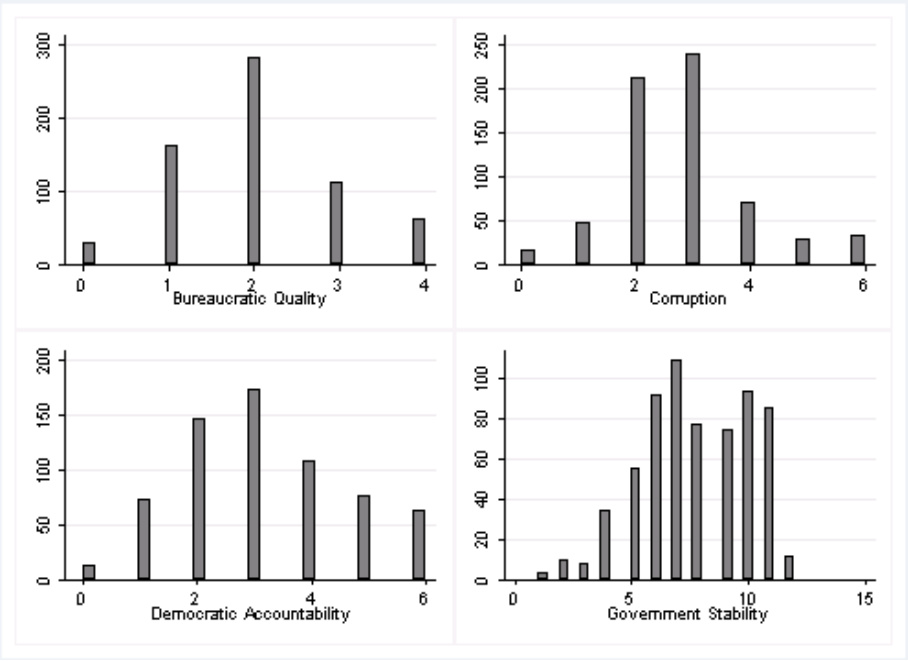

Figure A1: Distributions of ordinal variables, Bureaucratic Quality, Corruption, Dempcratic Accountability and Government Stability.

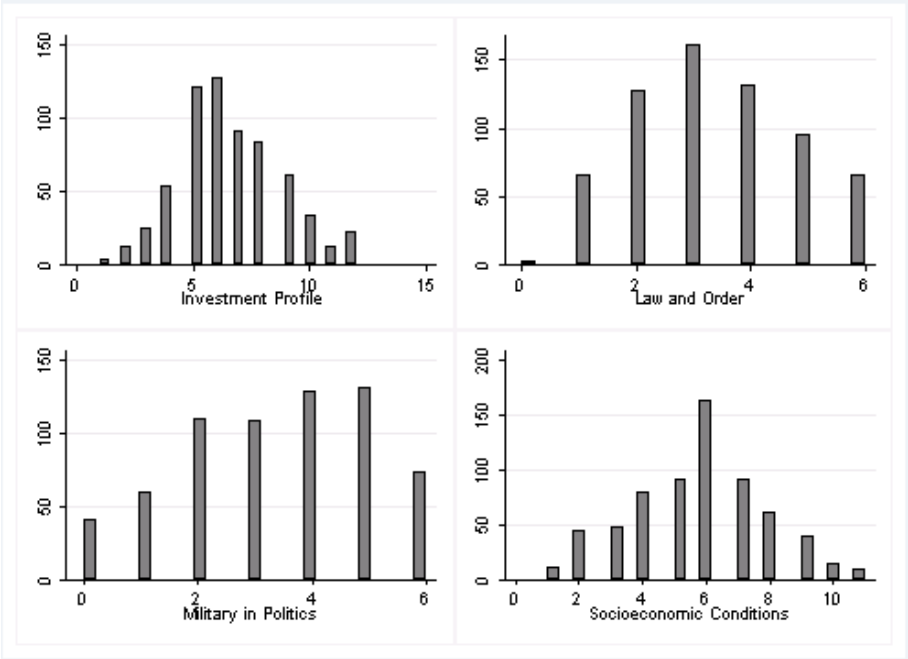

Figure A2: Distributions of ordinal variables, Investment Profile, Law and Order, Military in Politics and Socioeconomic Conditions.

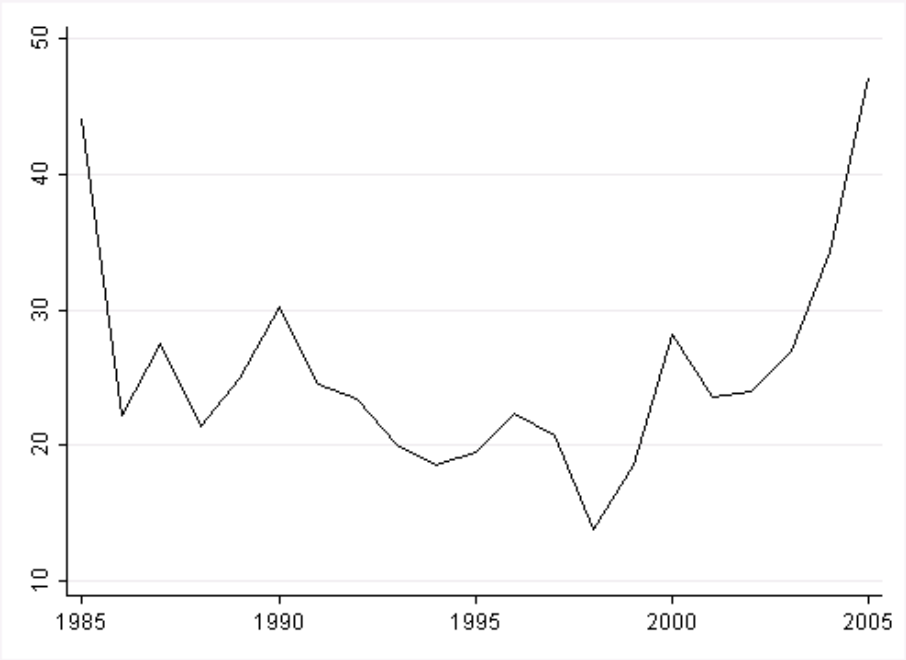

Figure A3: Oil price in year 2000 US dollar, 1985-2005.

Table A3: Robustness to un-interacted oil price and institutional variables.

Productivity Differential	0.680***
	(.142)
Ratio of Oil Exports to GDP	0.100*
	(.058)
Oil Price	0.001
	(.003)
Institutional variables	
Bureaucratic Quality	-0.571**
	(.028)
Corruption	-0.055
	(.035)
Democratic Accountability	0.089***
	(.024)
Investment Profile	-0.024
	(.016)
Law and Order	0.095***
	(.030)
Low Government Stability	-0.018
	(.013)
High Government Stability	-0.001
	(.040)
High Military Presence in Politics	0.001
	(.022)
Low Military Presence in Politics	0.007
	(.031)
Bad Socioeconomic Conditions	-0.02
	(.031)
Good Socioeconomic Conditions	-0.034
	(.022)
Institutional variables interacted with the ratio of oil exports to GDP	
Bureaucratic Quality	-0.17*
	(.068)
Corruption	0.100
	(.058)
Democratic Accountability	-0.111***
	(.049)
Investment Profile	0.002
	(.026)
Law and Order	-0.255***
	(0.055)
Low Government Stability	-0.007
	(.030)
High Government Stability	-0.062***
	(.022)
High Military Presence in Politics	-0.019
	(0.022)
Low Military Presence in Politics	-0.043
	(.030)
Bad Socioeconomic Conditions	0.098***
	(.038)
Good Socioeconomic Conditions	0.067
	(.048)
R-squared	0.22
N	551

Table A4: Robustness to regional exclusion: Western countries, Middle East, and Africa.

	Excluding United Kingdom, Norway, and Canada (1)	Excluding Egypt, Iran, Kuwait, Oman, Saudi Arabia, Yemen (2)	Excluding Algeria, Cameroon, Congo (Brazzaville), Congo (Kinshasa), Gabon, Libya, and Nigeria (3)
Productivity Differential	0.720***	0.525***	0.481***
	(.129)	(.178)	(.139)
Ratio of Oil Exports to GDP	0.122***	0.156***	0.085*
	(.045)	(.041)	(.042)
Institutional variables interacted with the ratio of oil exports to GDP			
Bureaucratic Quality	-0.163***	-0.139*	-0.192***
	(.045)	(.064)	(.033)
Law and Order	-0.182***	-0.127***	-0.109***
	(.040)	(.055)	(.037)
High Government Stability	-0.061***	-0.072	-0.078
	(.023)	(.049)	(.053)
High Military Presence in Politics	-0.043**	0.038	-0.116***
	(.022)	(.022)	(.021)
Bad Socioeconomic Conditions	0.072***	0.042	0.061***
	(.026)	(.033)	(.022)
R-squared	0.22	0.11	0.19
N	488	452	421

Notes: The dependent variable is the combined time series of the real effective exchange rate and bilateral real exchange rate index variables in logs. Institutional variables are measured so that higher values indicate stronger/better institutions, and are standardized with mean zero and standard deviation one. The estimation includes country-specific fixed effects, and Driscoll-Kraay standard errors are reported in parenthesis: *, significant at 10%; **, significant at 5%, *** significant at 1%.

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