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# Does deeper integration enhance spatial advantages? Market access and wage growth in China ☆

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ABSTRACT

New economic geography models predict that costly transport and the spatial distribution of demand affect the profits firms can earn in different locations, leading to higher wages for workers employed in cities with better geographic access to markets. In light of the dramatic embrace of globalization and labor market reforms that occurred in China after 1995, we measure the extent to which the influence of market access on wages strengthened and influenced wage growth over the subsequent period. Using survey data from two waves of the Chinese Household Income Project, we find that urban wages became more strongly influenced by access to markets, including domestic markets, between 1995 and 2002. The estimated elasticity of the wage with respect to market access of the worker's location more than doubles over the period. We also find that market access influences wages paid to both skilled and unskilled workers. Within provinces, we find no significant relationship between market access and either group's wages when adjusted for living costs, as expected in the context of internal labor migration. However, across provinces wages net of living costs are positively correlated with the market access of the worker's location. Consistent with deregulation of wage setting in state enterprises, the influence of market access on wages strengthened most for state-owned firms. A decomposition of the change in the mean wage indicates that market access is an economically important factor explaining the growth in average wages between 1995 and 2002.


## 1. Introduction

China's rapid integration into the global economy has reordered the international division of labor. Within China, it has helped to raise wages, while also increasing domestic income inequality. Over the past 25 years, China's real gross domestic product has expanded at an average rate of 9% per year. Much of this income growth has been driven by export production, with the growth in foreign trade averaging 15% annually since 1978. Early in the reform period, China integrated with the world economy only at its margins, allowing market sales of goods produced in excess of those meeting the planning quota, opening specific cities and zones to foreign trade and investment, and authorizing only a limited number of firms to engage in foreign trade. After 1992, China's international engagement broadened, as it permitted a surge in foreign investment, reduced tariffs, and expanded the number of authorized trading companies. As Branstetter and Lardy (2008, p. 633) stress, however, China's drive to liberalize trade and foreign direct investment dramatically accelerated in the late 1990s. Paving the way for eventual accession to the World Trade Organization in 2001, China unilaterally reduced tariffs on imports, dramatically cut quantitative restrictions on imports, eliminated many restrictions on foreign direct investment, and expanded public investment in roads, ports, airports and communications capacity.

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Perhaps not as well recognized as China's accelerated integration into global markets is the depth of its labor market reforms during the same period. In 1994, China passed the Labor Law, which took effect in 1995.[1] This law established the legal framework for worker–employer relations in the context of expanded employment flexibility. In addition to providing a framework for safeguarding workers' rights, the Law calls for equal treatment of workers across ownership sectors and permits no-fault dismissal of workers. Toward the end of the decade and in recognition of the employment mobility unleashed by the Labor Law, the Chinese government began strengthening social insurance programs and improving conditions for the increasingly large number of migrant workers. Although far from fully realized, these regulatory and policy changes represent significant steps toward market-mediated labor relations.

The combination of deeper international economic integration and labor market liberalization undertaken by China in the second half of the 1990s clearly influenced domestic wages and employment. Cai, Park, and Zhao (2008) report that mean annual real wages increased at an annual rate of 14% from 1997 to 2003.[2] Freeing firms to set wages according to market forces also led to increased spatial wage dispersion. Knight and Song (2008, p. 236), through an analysis of data from household surveys, find that about 11% of urban wage inequality is explained by the province in which the worker is employed.[3]

This paper investigates the spatial dimension of urban wages in China, adopting the new economic geography (NEG) approach taken by Hering and Poncet (2010). Our goal is to link cross-city variation in wages to spatial differences in access to markets. NEG models predict that costly transport and the spatial distribution of demand affect the prospective profits firms can earn in alternative locations, leading to the concentration of production in cities with better access to markets. If factor supplies are sufficiently different across locations so that factor prices are not equalized, workers in cities that offer firms higher potential profits, because of better "market access," will receive higher wages, specifically higher wages before adjustment for spatial cost-of-living differences. Hering and Poncet estimate Mincer (1974) equations using household data from 1995 and find that Chinese inter-city differences in individual wages can be partly explained by variation in the market access of the worker's location. They also find that the relationship is stronger for skilled workers and those employed by foreign-invested firms.

In light of the dramatic embrace of globalization and labor market liberalization that occurred in China after 1995, we revisit and extend the work of Hering and Poncet (2010). First, using household survey data from the 1995 and 2002 waves of the Chinese Household Income Project (CHIP), we measure the extent to which the influence of market access on wages changed over this period of intense domestic labor market liberalization and international integration. We measure changes in the influence of market access on wages of different types of workers and for different types of firms. Secondly, we examine the relationship between market access and real wages, defined as market wages adjusted for the local cost of living, seeking to find evidence of the extent to which internal migration offset the spatial advantages. Lastly, we decompose the change in mean wages and estimate the extent to which income growth between 1995 and 2002 was related to stronger spatial advantages made possible by continuing integration into global markets.

We find that urban wages did become strongly influenced by access to markets, including domestic markets, between 1995 and 2002. The estimated elasticity of the wage with respect to market access of the worker's location more than doubles over the period. We also find that market access influences wages paid to both skilled and unskilled urban workers but that this relationship is significantly stronger by 2002 only for skilled workers. We use quantile regressions to explore the possibility that more able workers migrate to central locations and find that market access is a significant determinant of wages for all quantiles. Consistent with a reduction in barriers to within-province migration, we find no significant relationship between market access and worker's wages when we adjust them for cost-of-living differences and include provincial fixed effects in our estimating equation. These results suggest that, within provinces, higher nominal wages made possible by the cost advantages of central locations attract workers from other areas, transferring spatial advantages to fixed local factors, such as land and fixed housing stock. When we exclude provincial fixed effects from the Mincerian wage regressions, however, we find that workers in cities with better market access earn higher real wages. While we are unable to rule out the possibly confounding influence of unobserved provincial characteristics, this result suggests that inter-provincial migration is not sufficient to equalize workers' real rewards and that the growing importance of spatial advantages in factor rewards is another source of coastal-inland wage inequality.

We also find that the responsiveness of wages to market access depends on the employer's ownership type. While the estimated wage elasticity is lower for state-owned enterprises than for private or foreign firms, the relationship between market access and wages grew most quickly for state-owned firms, in keeping with the deregulation of wage setting in this sector. A decomposition of the change in the mean wage indicates that market access is an economically important factor explaining growth in average wages between 1995 and 2002. At least 7% of the difference over time is due to the growing wealth of markets close to Chinese producers, especially those within China itself. At least another 13% of wage growth is due to the stronger influence over time of market access on wages.

## 2. Theoretical framework and its application to China

### 2.1. Theoretical framework

NEG models emphasize the interplay of transport costs and plant-level increasing returns to scale. Beginning with Helpman and Krugman (1985) and Krugman (1991), the Dixit–Stiglitz model of monopolistic competition has been used to derive a

[1] Gallagher and Jiang (2002) provide an overview and analysis of the Labor Law of the People's Republic of China.

[2] Cai, Park, and Zhao (2008, p. 184) provide a time series for 1978 through 2003.

[3] Income inequality, as well as wage inequality, has a spatial dimension. Gustafsson et al. (2008, p. 54) report that in 2002 19% of urban income inequality can be attributed to provincial differences in mean incomes.

relationship between the distribution of expenditure across trading economies and the distribution of production. Plant-level scale economies imply that firms find it most profitable to concentrate production at one location, with that location chosen to minimize transportation costs to the firm's geographically dispersed markets. Redding and Venables (2004) show that the potential net profits earned in each location are a linear function of the location's "market access," which is essentially a discounted sum of local and foreign expenditures available to the firm.

Predictions of the model for production and factor prices depend on the full general equilibrium. If factor price equalization (FPE) obtains in equilibrium, Behrens, Lamorgese, Ottaviano and Tabuchi (2004) provide tests of the model appropriate for a multi-country world.[4] The model predicts that larger and more centrally located countries attract a disproportionate share of firms and account for a disproportionate share of production. Their approach has been applied to data from 57 European regions over the period 1985–2000 by Head and Mayer (2006), who find little support for the model characterized by FPE. More recently, Niepmann and Felbermayr (2010) investigate the implications of the Behrens et al. model for the response of production shares to decreasing trade costs. Using data from 20 OECD countries for 1980 to 1999, they find that the distribution of firms across countries becomes increasingly skewed as trade becomes freer, a finding consistent with the model's theoretical predictions.

Because these approaches rely on FPE, they are not appropriate for a study focused on regional disparities in China. As noted above, wages are not equalized across Chinese regions; studies using individual level data have found a strong relationship between a worker's location and his or her wage. An alternative approach is provided by Redding and Venables (2004), who focus on how market access influences wages in a world in which FPE does not obtain. Taking the size of markets and the number of competitors serving those markets as exogenous, the theory relates a location's market access to the wages manufacturing firms are able to pay there. In equilibrium, wages in places with better market access are higher so that profits are equalized across locations and firms have no unilateral incentive to move.

The Redding and Venables (2004) framework leads to a log–linear relationship between the wage paid in a given location and a discounted sum of the economic activities in all relevant regions accessible from this location.[5] The NEG prediction is that wages will be higher in locations with better market access, or that the estimated coefficient for a market access measure in a wage regression will be positive and statistically significant. Redding and Venables apply this approach to an analysis of variation in GDP per capita across 110 countries, from 1992 to 1996, and find evidence consistent with the theory. Similarly, Head and Mayer (2006) investigate the dispersion of average wages across European regions and obtain results consistent with the NEG prediction.

A significant innovation in Redding and Venables (2004) is the use of a gravity model to measure "real market potential," as Head and Mayer (2006) term the Redding–Venables market access term, and also to estimate the effect of physical distance on trade. The NEG model provides for a formal definition of real market potential, showing its dependence on expenditures in each potential market, but also on the capabilities of potential competitors at these locations. This real market potential measure is often distinguished from the Harris (1954) market potential measure, which implicitly treats competitor capabilities as constant across all location and, thus, is often considered a "nominal market potential" measure. Redding and Venables show how exporter fixed effects recovered from a gravity model of international trade can be used to measure real market potential.

In their analysis of 1995 Chinese urban wages, Hering and Poncet (2010) make an important technical contribution by combining an analysis of individual, rather than average, wage levels with the Redding–Venables measure of market access.[6] To create this measure, they use geographically detailed Chinese trade data to estimate a gravity model and recover the importer and exporter fixed effects. Their procedure entails the use of bilateral trade data at the provincial level merged with trade flows from other countries in the world trading system. Further disaggregation of market potential to the city level requires an allocation of provincial market access among cities based on city shares of provincial GDP. Their procedure requires use of restricted China Customs data recording import and export values to or from each province to all international partners. Controlling for individual characteristics, province and sector of employment, they estimate a positive and statistically significant relationship between a city's market access and Chinese urban workers' wages.

### *2.2. Liberalization and the strength of the market access effect, 1995 to 2002*

While Hering and Poncet (2010) provide evidence that the size of foreign markets influenced Chinese wages as early as 1995, the purpose of this paper is to measure the extent to which wages reflect market access in 2002, and also the contribution of market access to growth in average urban wages between 1995 and 2002. Given extensive changes in Chinese commercial policy and labor market regulation over the period, we predict that market access has become a more important determinant of spatial wage inequality. Our prediction rests on two observations about policy changes in China occurring over this horizon. First, liberalization of China's commercial policy and expansions in domestic infrastructure substantially reduced trade frictions, strengthening the relationship between domestic factor prices and foreign markets. Secondly, even if no changes to commercial policy had been made over the period, changes in Chinese labor market regulation allowed wages to increasingly reflect market determinants. We briefly review changes in Chinese commercial policy and labor market regulation between 1995 and 2002 to support these claims.

[4] Behrens, Lamorgese, Ottaviano and Tabuchi (2009) provide predictions for the case in which FPE does not hold.

[5] Fujita, Krugman, and Venables (1999) provide a fully specified general equilibrium model in which this relationship between market access and wages is identified.

[6] In a recent paper, Fally, Paillacar, and Terra (2010) also estimate the impact of market access on wages controlling for individual worker characteristics. Their study uses Brazilian microdata and also controls for supplier access and firm productivity. They find a positive and significant effect of market access on measured differences in wages across Brazilian states and industries.

Branstetter and Lardy (2008) describe in detail China's changing commercial policy in the second half of the 1990s. As noted above, China undertook significant unilateral trade liberalization in anticipation of WTO accession (p. 634–9). The government also made far reaching changes to its FDI regime that greatly expanded the freedom with which foreign firms could operate in China (p. 645). Branstetter and Lardy conclude that "the view of the export sector as an enclave with little connection to the local economy became increasingly out of date by the eve of China's accession to the WTO (p. 638)".

Branstetter and Lardy (2008) also question the depiction of China's internal markets as deeply fragmented. While a number of studies document the extent of interprovincial protectionism and market fragmentation (Poncet 2003, Poncet, 2005, Wedeman, 2003, Young, 2000), other research suggests that internal markets are substantially integrated. Bai, et al. (2004) and Holz (2009) provide extensive analysis of industrial specialization, interprovincial trade, and inter-city price dispersion to contradict this characterization. Branstetter and Lardy weigh the conflicting evidence and conclude that "international integration appears to be proceeding along with intranational integration." With this view, access to domestic, as well as foreign, markets could be viewed as an increasingly important determinant of urban wage levels.

While adjusting to the extensive commercial policy reforms undertaken after 1995, the Chinese labor market also was subject to seismic changes unleashed by regulatory reform. As noted by Cai, Park, and Zhao (2008), the 1994 Labor Law facilitated the massive restructuring of state-owned enterprises, which led to layoffs of at least 10 million workers by 1997 and 27 million more from 1998 to 2004.[7] Giles, Park, and Zhang (2005) estimate that the unemployment rate for all urban residents rose from 6.8% in 1996 to 11.1% in 2002. Workers remaining in the state sector were subject to removal of most administrative controls on the determination of wages and benefits. Thus, over the period, wages were largely freed to adjust to market-determined levels.

Hering and Poncet (2010) characterize 1995 as a time in which internal migration was unlikely to affect urban wages.[8] Between 1995 and 2002, however, central and local governments loosened controls on domestic migration, including migration without formal changes in household registration (non-*hukou* migration). Cai, Park, and Zhao use the 2000 census to calculate the stock of migrants, defined as the share of persons residing in a location for more than six months in the prior year whose *hukou* is from outside the city or county. By this definition, they estimate that by the turn of the century migrants comprised 14.6% of the population and 19.6% of the employment in China's cities (p. 191). Therefore, by 2002 labor markets, as well as product markets, were significantly freer and more open than they had been in 1995.

To summarize, we test the NEG model prediction that variation in wages reflects variation in access to domestic and foreign markets. We adapt the approach taken by Hering and Poncet (2010) for 1995 and extend it to an analysis of individual urban wages in 2002, and compare across the two sample years. We predict that (a) wages in both periods are positively and significantly related to the market access of the city in which the worker is employed; and (b) this relationship became significantly larger by 2002. We estimate the relationship for all urban workers, for skilled urban workers and for unskilled urban workers. We also test for differences in the strength of the effect by firm ownership type, state-owned, private-owned, and foreign-owned.

## 3. Econometric specification, data, and measurement

### 3.1. Econometric specification

To test our predictions, we specify augmented Mincer equations for 1995 and 2002, estimating the influence of market access on wages in each location. We then test whether the strength of this relationship is statistically different across the two years. We apply this relationship to data on the hourly wages, $w$, of individual Chinese urban workers, taking the market access ($MA$) of region $j$ as exogenous. In its simplest form, letting $X_i$ be a vector of personal characteristics of worker $i$ who is employed in region $j$, this relationship can be expressed as:

$$\ln w_{ij} = a + b \ln MA_j + \gamma X_i + \varepsilon_{ij}, \tag{1}$$

where $a$ and $b$ are (scalar) coefficients to be estimated, $\gamma$ is a vector of coefficients to be estimated, and $\varepsilon_{ij}$ is an error term assumed to be identically and independently distributed. We estimate Eq. (1) using OLS, allowing for constants specific to province and sector of employment, and in other specification, for constant specific to each province-sector. Observable and unobservable characteristics of workers in the same location may be correlated, causing OLS measure to understate true standard errors. We correct for this problem by clustering errors at the sector-province level and report only robust standard errors.

To see if differences in the market access effect are stronger in 2002 than in 1995, we estimate (1), pooling the data across years and interacting all terms with a year indicator variable. This method allows us to test whether cross-year differences in the estimated coefficients are statistically significant. Next, separating the sample into skilled and unskilled workers on the basis of their educational attainment, we see if our findings hold for both groups of workers. We also interact the market access measure with the employer's ownership type and test whether market access is a significant determinant of wages for state-owned firms as well as private sector employers.

Our data also allow us to explore whether real wages are related to spatial characteristics of the worker's location. When workers are free to migrate from one location to another, spatial equilibrium requires that workers as well as firms have no unilateral incentive

[7] See Cai, Park, and Zhao (2008) for details of SOE restructuring and analysis of labor market responses.

[8] See Hering and Poncet (2010), footnote 7, where they argue that the impact of migrants on urban wages is negligible since urban labor markets are strictly segmented. Our point is that this segmentation deteriorated between 1995 and 2002 due to a relaxation of formal controls.

to change location. Hanson (2005) follows Helpman (1998) and posits that areas with high market access will have high nominal wages to keep firms in equilibrium while prices for local nontradeables, such as housing, will rise to equalize real wages. We test this prediction by replacing the market wage on the left-hand side of Eq. (1) with the real wage, defined as the market wage adjusted for living costs. If migration is sufficient to equalize real wages across locations, there should be no statistically significant relationship between market access, *MA*, and the real wage. We can also test whether this relationship was the same in 1995 and 2002.

Lastly, we test whether the relationship between measured market access and wages is robust to the inclusion of controls for local agglomeration of economic activity. We note that NEG theory predicts that spatial differences in market access will lead to industrial concentration, and thus agglomeration can be viewed as an outcome of the same process that leads to higher wages in spatially advantaged cities. However, agglomeration can be the outcome of other processes that also raise wages and so we test whether the market access effect is robust to the inclusion of these controls.

### 3.2. Data description

To estimate the Mincer equations, we use data drawn from individual surveys of the Chinese Household Income Project (CHIP) for the years 1995 and 2002.[9] The 1995 surveys cover 6931 households and 21,698 individuals across 11 provinces; the 2002 surveys cover 6835 households and 20,632 individuals in 12 provinces.[10] The surveys do not follow the same individuals over time so CHIP is not an individual panel dataset. We restrict our attention to the sample of the urban population employed in nonagricultural tradable good sectors.[11] We further restrict the sample to only include workers who are between the ages of sixteen and sixty, not self-employed in their primary job and not missing information for years of schooling and years of work experience.[12] To control for the influence of extreme observations of our dependent variable, log hourly wage, we exclude outlier observations identified using Hadi's (1994) method.[13]

The dependent variable, log hourly wage, is constructed from annual deflated labor income and total annual work hours information. Nominal annual income in 2002 is converted to real 1995 yuan using province level Consumer Price Index (CPI) obtained from various issues of the China Statistical Yearbook.[14] Annual total labor income in 2002 is calculated as the sum of basic salary, bonus, allowances and subsidies, and living expenses for the laid-off.[15] Annual work hours are calculated by multiplying average work hours per day by average work days per month and finally by months worked. The reported average work time excludes periods of unemployment as well as days worked in informal activity while unemployed.

The CHIP data allows us to control for a wide set of individual characteristics, including gender, years of schooling, experience, age, communist party membership. These characteristics have been shown to be significant determinants of urban wages by Knight and Song (2008) in both 1995 and 2002.[16] We also control for the ownership type of the worker's employer, private enterprise, state enterprise, or foreign-invested enterprise. Table 1 contains a complete list of variables used and summary statistics for the sample of individuals included in our analysis.

### 3.3. Measuring market access

While we would prefer to measure market access using the Redding–Venables method, we do not have access to the proprietary China Customs data used by Hering and Poncet (2010) for their application of the method to Chinese provincial trade. As an alternative, we use a measure of market access that is similar to the Harris (1954) nominal market potential measure, with the distance discount factor drawn from Hering and Poncet's gravity model. In other words, we implicitly assume that the competitive conditions China faces in each potential market is constant across destinations. We do not consider our use of nominal market potential, rather than the Redding–Venables real market potential, to be a major drawback to our analysis for several reasons. First, our study concerns variation across Chinese cities. It is clear that the only difference in access to foreign markets between any two Chinese cities is the distance to the nearest port. It is these differences, rather than the market size or competitive

[9] 1995 surveys accessed at http://www.icpsr.umich.edu/icpsrweb/ICPSR/studies/03012; 2002 surveys accessed at http://www.icpsr.umich.edu/icpsrweb/ICPSR/studies/21741. We thank Laura Hering and Sandra Poncet for sharing the final sample of the 1995 CHIP that was used in Hering and Poncet (2010).

[10] The data appendix identifies the cities included in both the 1995 and the 2002 CHIP urban surveys.

[11] Hukou status is provided for the 2002 sample only and in that sample 99% of respondents have urban hukou in the resident city.

[12] The questionnaires ask for employment information in the primary and any secondary jobs. Workers are classified based on their primary job situations since most people do not report having second jobs.

[13] See Hering and Poncet (2010) for details on the construction of the final 1995 sample. Of the 6079 individuals in the final sample, we could not construct the market access variable for 555 individuals due to missing information on GDP or land area for seven cities of Fenyang, Bozhou, Tianmen, Geiju, Dali, Wuwei, and Pingliang. Therefore, the final 1995 sample consists of 5555 individuals. In the original 2002 sample 4609 individuals work in the three nonagricultural tradable sectors. Of these, we could not identify the exact location of 76 individuals, 629 are self-employed and 27 are not between the ages of sixteen and sixty. A further 74 are identified as outliers via Hadi's (1994) method. Of the 3803 remaining individuals, 10 have missing wage information. The market access variable could not be constructed for four cities of Fenyang, Geiju, Pu'er Hani-Yi and Dali due to missing information on GDP or land area which led to a further reduction of 135 individuals living in these cities. Therefore, the final 2002 sample consists of 3658 individuals.

[14] The China Statistical Yearbooks report province level CPI with the base year set to the preceding year. The CPI was re-indexed setting 1995 as the base year.

[15] We exclude income not directly from the primary employer, subsidy for minimum living standard, living hardship subsidies from work unit, the second job and sideline income and monetary value of income in kind. We recalculate hourly wage based on all the above sources of income and find the estimation results to be qualitatively unchanged.

[16] Knight and Song (2008) use a slightly different set of individual characteristics to explain the variation in urban wages. We follow the specification used by Hering and Poncet (2010) to facilitate comparison with their results. We note that the exact set of individual controls included in the estimation does not affect our qualitative results.

**Table 1**
Variable definitions, sources, and summary statistics.

| | | | 1995 | | 2002 | |
|---|---|---|---|---|---|---|
| Variable | Definition | Source | Mean | St. Dev. | Mean | St. Dev |
| Log hourly wage | Annual labor income divided by the total annual work hours, in 1995 yuan | Urban Individual Data, 1995 and 2002 | 0.88 | 0.66 | 2.03 | 0.82 |
| Female | Categorical variable equal to 1 if individual is female | Urban Individual Data, 1995 and 2002 | 0.50 | 0.50 | 0.43 | 0.49 |
| Years of schooling | Number of years of education | Urban Individual Data, 1995 and 2002 | 10.06 | 2.81 | 10.77 | 2.68 |
| Experience | Number of years of work experience | Urban Individual Data, 1995 and 2002 | 19.91 | 9.64 | 20.82 | 9.35 |
| Age | Age of individual | Urban Individual Data, 1995 and 2002 | 39.14 | 9.99 | 40.54 | 8.85 |
| Age$^2$ | Age squared | Urban Individual Data, 1995 and 2002 | 1631.71 | 787.55 | 1721.41 | 702.92 |
| Communist | Categorical variable equal to 1 if individual is a communist party member | Urban Individual Data, 1995 and 2002 | 0.19 | 0.39 | 0.24 | 0.42 |
| Log minimum living cost | The average household expenses per month needed to maintain a minimum living standard in the city where the worker resides, in 1995 yuan | Urban Household Data, 1995 and 2002 | 6.65 | 0.30 | 7.70 | 0.35 |
| Log market access (MA) | Sum of international and domestic market access, in billion 1995 yuan[a] | China Statistical Yearbook, 1998 and 2003; Urban Statistical Yearbook, 1998 and 2003 | 0.38 | 0.40 | 0.67 | 0.46 |
| Log employment density | Total employment in a city, per km$^2$ | 1998 and 2002 Annual Surveys of Industrial Production | 3.42 | 1.08 | 2.97 | 1.22 |
| Log firm density | Total number of firms in a city, per km$^2$ | 1998 and 2002 Annual Surveys of Industrial Production | −2.57 | 1.07 | −2.77 | 1.26 |
| Log population | Total population in a city, in ten thousand people | Urban Statistical Yearbook, 1998 and 2003 | 6.16 | 0.51 | 6.33 | 0.61 |

Notes: The table reports the sample means. There are 44 and 56 cities in 1995 and 2002 respectively. Domestic market access is the city-specific, distance-weighted sum of Chinese city 1997 and 2002 GDP, in billion 1995 yuan.
[a] International market access is the city-specific, distance-weighted sum of export partner 1997 and 2002 GDP, in billion 1995 yuan.

conditions of foreign markets that drive the cross-city variation in market access that we use to identify the market access effect. Secondly, a direct comparison of estimation results using the Redding–Venables measure and the Harris measure is provided by Head and Mayer (2006). They report that while the Harris measure yields a larger estimated market access coefficient, the root mean squared error of the regressions are virtually identical. Either measure, as well as other alternatives, produces a significant market access coefficient.

To create our market access measure, we decompose total market access of a city into domestic and international market access:

$$MA_c = MA_c(D) + MA_c(I), \tag{2}$$

where $c, D, I$ denote city, domestic, and international respectively. Each component is a distance weighted sum of real GDP as described below.

### 3.4. Domestic market access

Domestic market access measures the potential demand faced by city $c$ firms from other cities within China. It can be thought of as the sum of demand from within the city itself, from other cities within the province it resides in, and from all other cities in China.

$$MA_c(D) = GDP_c/dist_{cc}^{\delta} + \sum_{k \epsilon Province} GDP_{c_k}/dist_{cc_k}^{\delta} + \sum_{l \epsilon China} GDP_{c_l}/dist_{cc_l}^{\delta} \tag{3}$$

where the first term is the city's own GDP weighted by $dist_{cc}^{\delta}$, the distance between city $c$ and itself, and $\delta$ is a discount factor.[17] We assume the discount factor to be 1.5 in all our regressions, drawing this from the gravity estimates for Chinese provinces in Hering and Poncet (2010).[18] The second term is the sum of GDP of all other cities, $c_k$, in province $k$ where city $c$ is located but excluding city $c$, weighted by the sum of distance between city $c$ and all other cities within province $k$. The final term is the sum of GDP of all other cities, $c_l$, in provinces excluding province $k$, weighted by the sum of distance between city $c$ and these cities. We use city level GDP data in 1997 and 2002 to construct (3).[19] City level GDP is converted to real values using the national GDP deflator.[20] Distance, in kilometers, is calculated as the arc distance between the geographic centroid of each city using ArcGIS software.

[17] The distance between a city and itself is calculated as the average distance between a producer and consumer in a circular country as $\left(\frac{2}{3}\sqrt{area_c/\pi}\right)^{\delta}$ following Amiti and Javorcik (2008), Head and Mayer (2004), Hering and Poncet (2010), and Redding and Venables (2004).
[18] We carried out robustness checks using a range of values between 1 and 2 and we find our benchmark results to remain qualitatively unchanged.
[19] City level GDP data is unavailable prior to 1996. We follow Hering and Poncet (2010) who construct their market access variable using information as of 1997.
[20] We also deflate city GDP using province level CPI and the results are qualitatively unchanged.

### 3.5. International market access

International market access measures the potential demand faced by city *c* firms from the rest of the world:

$$MA_c(I) = \sum_{j\ Export\ Partners} GDP_j / \left(dist_{cp} + dist_{pj}\right)^{\delta}, \tag{4}$$

where *j* and *p* denote export partner and port city respectively. International market access is calculated as the sum of GDP of China's 40 major export destinations weighted by the sum of the distance between city *c* and the nearest port city *p* and the distance between the port city and the capital city of the partner country, *j*.[21] Export partner GDP is drawn from the World Bank's World Development Indicators.[22] There are three major port cities in China – Guangzhou, Shanghai, and Tianjin. We assume that goods are first delivered from a city to a port prior to being exported out of China (Lin, 2005). Therefore, we first determine which of these three port cities is closest to city *c* and then calculate the arc distance between the two cities and finally the arc distance between the closest port city and the capital city of China's export destinations.[23]

Fig. 1 shows total market access and its constituent parts for all cities that appear in both waves of the CHIP. It is interesting to note that there is very little growth in the average distance weighted foreign real market access, MA(I), as shown by the first column for each year. In contrast, due to the rapid growth of real GDP within China, the average distance-weighted domestic market access, MA(D), almost doubles, as shown by the second columns in Fig. 1. It is domestic market growth that drives the 75% increase in the value of the average market access measures for cities in our sample.

## 4. Estimation results

### 4.1. Market access and individual wages, 1995 and 2002

Results of estimating the augmented Mincer Eq. (1) for the full sample of urban workers are shown in Table 2. The first panel provides wage regressions for 1995 and 2002 estimated without including the market access measure. These equations also contain occupation controls and sector-province fixed effects. We cluster errors at the sector-province level. As expected from previous research, all of the individual worker characteristics, as well as indicators for employer ownership type, are highly statistically significant and have the expected sign. Overall, the included variables explain 52% of the variance in wages. The third column of the panel provides the difference in the coefficient estimate for 2002, relative to that estimated for 1995. We find that the penalty for being female is somewhat higher in 2002, experience is somewhat less valued, and the return to education has doubled. Knight and Song (2008) use the same dataset to provide an extensive discussion and analysis of changes over time in urban wage determinants.

Our focus here is on the influence of market access on urban wages and coefficient estimates for 1995 and 2002 can be found in panels (2) and (3) in Table 2. Panel (2), estimated with sector fixed effects and province fixed effects, and panel (3), estimated with sector-province fixed effects, tell the same story. The estimated elasticity of wages with respect to the market access of the worker's location is 0.28 for the 1995 sample and 0.67 for the 2002 sample, and both are highly statistically significant. As predicted from our consideration of commercial and labor market policy changes, wages are more responsive to spatial differences in potential profits by 2002. The estimated elasticity for 2002 is 2.4 times the elasticity estimated for 1995 and this difference is highly significant. Perhaps surprisingly, the fit of the equation is not noticeably improved by the inclusion of our market access measure, which may be due to the inclusion of province fixed effects in all equations as these control for a large share of the variation in access to coastal ports and domestic markets.[24]

### 4.2. Quantile regression

In our basic estimations, shown in Table 2, we achieved an R-squared value of 0.52, indicating that the explanatory variables in the base model explain only about 50% of the variation in individual hourly wages. Although we control for an individual worker's years of schooling, we do not have an individual level panel and, therefore, cannot include individual fixed effects to control for individual unobservables. A particular concern is that more able workers have a higher likelihood of migrating to larger cities.[25]

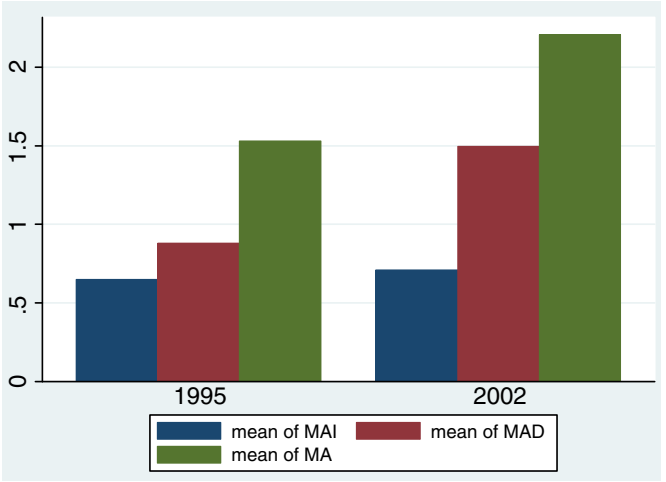
[21] The countries considered are Argentina, Australia, Austria, Belgium, Brazil, Brunei, Cambodia, Canada, Chile, South Korea, Denmark, Egypt, Finland, France, Germany, Hong Kong, India, Indonesia, Italy, Japan, Laos, Macao, Malaysia, Mexico, Netherlands, New Zealand, Philippines, Poland, Russia, Saudi Arabia, Singapore, South Africa, Spain, Switzerland, Taiwan, Thailand, U.A.E., U.K., U.S.A., and Vietnam. These countries together represented approximately 92% and 90% of China's total exports in 1997 and 2002 respectively.

[22] GDP in constant 2000 USD was converted to 1995 USD using the implicit GDP price deflator in Table B-3, *Economic Report of the President*, U.S. Government Printing Office, 2006. It was converted to Chinese RMB by applying the 1995 official exchange rate, 8.3514, published in the IMF's 2006 *International Financial Statistics*.

[23] To calculate the distance between a port city and capital city of China's export partners we first obtain the latitude and longitude points for each location at http://www.mashupsoft.com/maps/latlonlocator. We input this information into the arc distance formula provided at www.nau.edu/~cvm/latlongdist.html to obtain the arc distance between any pair of locations.

[24] While it would be desirable to decompose the market access measure and allow separate coefficients for domestic and international market access, the two measures have a very high cross-sectional correlation (0.83) and this makes identification of the separate influences difficult. Unreported regression results suggest that the influence of domestic MA has risen over time, consistent with decreased interprovincial trade barriers.

[25] A higher migration probability has been estimated for more highly educated workers by Knight and Song (1999) using a rural household survey and logistic regression.




Notes: MA refers to total market access, MAI refers to international market access, and MAD refers to domestic market access, as defined in the text. The graph shows the mean value of each measure for the sample of cities that appears in both the 1995 and 2002 waves of the Chinese Household Income Project.

**Fig. 1.** Composition of the market access measure (in 1995 billion yuan), 1995 and 2002.

**Table 2**
Market access and wages, 1995 and 2002.

| Dependent variable: log hourly wage | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | | | (2) | | | (3) | | |
| | 1995 | 2002 | Difference | 1995 | 2002 | Difference | 1995 | 2002 | Difference |
| *Market access* | | | | | | | | | |
| Log MA | – | – | – | 0.281*** | 0.666*** | 0.385*** | 0.286*** | 0.668*** | 0.382*** |
| | | | | (0.034) | (0.067) | (0.067) | (0.041) | (0.065) | (0.069) |
| *Individual characteristics* | | | | | | | | | |
| Female | −0.100*** | −0.149*** | −0.049** | −0.098*** | −0.141*** | −0.042 | −0.099*** | −0.138*** | −0.040* |
| | (0.017) | (0.019) | (0.023) | (0.021) | (0.019) | (0.034) | (0.018) | (0.020) | (0.023) |
| Years of schooling | 0.023*** | 0.046*** | 0.023*** | 0.023*** | 0.043*** | 0.021*** | 0.022*** | 0.042*** | 0.020*** |
| | (0.004) | (0.005) | (0.006) | (0.002) | (0.004) | (0.003) | (0.004) | (0.005) | (0.006) |
| Experience | 0.017*** | 0.005* | −0.012*** | 0.017*** | 0.004 | −0.013*** | 0.017*** | 0.004 | −0.013*** |
| | (0.002) | (0.003) | (0.004) | (0.002) | (0.005) | (0.004) | (0.002) | (0.003) | (0.004) |
| Age | 0.068*** | 0.062*** | −0.006 | 0.067*** | 0.070*** | 0.004 | 0.067*** | 0.068*** | 0.001*** |
| | (0.011) | (0.013) | (0.015) | (0.007) | (0.013) | (0.006) | (0.011) | (0.013) | (0.000) |
| $Age^2$ | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Communist | 0.073*** | 0.079** | 0.006 | 0.077*** | 0.092** | 0.016 | 0.075*** | 0.086** | 0.011 |
| | (0.024) | (0.029) | (0.036) | (0.016) | (0.034) | (0.032) | (0.024) | (0.029) | (0.036) |
| *Ownership dummies* | | | | | | | | | |
| State-owned | 0.156*** | 0.143*** | −0.013 | 0.158*** | 0.148*** | −0.011 | 0.158*** | 0.141*** | −0.017 |
| | (0.025) | (0.039) | (0.034) | (0.011) | (0.023) | (0.034) | (0.024) | (0.037) | (0.034) |
| Foreign-owned | 0.326*** | 0.356*** | 0.030 | 0.316*** | 0.331*** | 0.015 | 0.316*** | 0.331*** | 0.015 |
| | (0.031) | (0.086) | (0.101) | (0.011) | (0.034) | (0.038) | (0.037) | (0.084) | (0.103) |
| Occupation dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector and province | | | | | | | | | |
| Fixed effects | No | No | No | Yes | Yes | Yes | No | No | No |
| Sector-province fixed effects | Yes | Yes | Yes | No | No | No | Yes | Yes | Yes |
| Observations | | 9213 | | | 9213 | | | 9213 | |
| $R^2$ | | 0.52 | | | 0.52 | | | 0.53 | |

Notes: Robust standard errors in parentheses. Private-owned is the left out category.
* 10% significance.
** 5% significance.
*** 1% significance.

**Table 3**
Market access and wages, quantile regression, 1995 and 2002.

| Dependent variable: log hourly wage | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Quantile | 0.125 | | | | 0.375 | | | 0.625 | | | 0.875 | |
| | 1995 | 2002 | Difference | 1995 | 2002 | Difference | 1995 | 2002 | Difference | 1995 | 2002 | Difference |
| *Market access* | | | | | | | | | | | | |
| Log MA | 0.355*** | 0.677*** | 0.324*** | 0.309*** | 0.622*** | 0.313*** | 0.221*** | 0.616*** | 0.395*** | 0.160** | 0.568*** | 0.408*** |
| | (0.061) | (0.070) | (0.093) | (0.045) | (0.048) | (0.066) | (0.048) | (0.052) | (0.071) | (0.060) | (0.068) | (0.091)* |
| Occupation dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector-province fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 9213 | | | | 9213 | | | 9213 | | | 9213 | |
| Pseudo $R^2$ | 0.33 | | | | 0.39 | | | 0.43 | | | 0.43 | |

Notes: Private-owned is the left out category. The specification is the same as in Table 2, panel 3; only the market access coefficients are reported.
* 10% significance.
** 5% significance.
*** 1% significance.

Because we cannot control for heterogeneous worker ability, we may erroneously attribute higher observed wages to greater local market access rather than to ability. To shed some light on what type of unobservable factors might be at play, we follow Knight and Song (2008) and estimate quantile regressions.

Quantile regression, introduced by Koenker and Bassett (1978), models the relationship between a set of explanatory variables, $X$, and the conditional quantiles of the outcome variable, $Y$ given $X=x$ in contrast to ordinary least-squares regression which models the relationship between $X$ and the conditional mean of $Y$ given $X=x$.[26] In the context of this study, this allows us to compare the wage equation across different points in the conditional wage distribution. Differences in the coefficients on an explanatory variable of interest across various quantiles could reveal a relationship between itself and the unobserved variables. We carry out quantile regressions for values of the quantile $q=0.125$, 0.375, 0.625, and 0.875 and the results are reported in Table 3.

The specification is identical to that in Table 2, but our focus is on the market access variable. We observe that in each year and at each quantile, the coefficient on market access is highly significant and similar across each of our four chosen quantiles. This is in contrast to what we would expect under our hypothesis that higher ability individuals tend to migrate to larger cities. Assuming that the conditional wages represent unobserved ability, we would expect the effect of market access to be lower for higher ability individuals and therefore decrease as we move up the quantiles. However, the estimated market access coefficient remains positive and highly significant and is not significantly different across the quantiles for each of the years, lending confidence that the coefficient on market access is not contaminated by the effect of unobserved worker ability.

### 4.3. Differences across education groups

Table 4 presents Mincer regressions including the market access measure, with the full sample split into two subgroups based on educational attainment. Skilled workers are defined as those who have completed at least nine years of schooling, while those with fewer years of schooling are classified as unskilled workers. For skilled workers, we see that market access is a significant determinant of the wage in both 1995 and 2002. As estimated using the full sample, the strength of the market access effect increased over the period and this difference is highly statistically significant. Surprisingly, perhaps, we estimate very similar wage elasticities for the unskilled worker sample, with the estimate for 2002 very close to that found for the skilled worker sample. However, there is more variation within this smaller subsample and the difference in the estimated coefficients across the years is not significant. Thus, while we estimate a larger elasticity for low skilled wages in 2002 than in 1995, we cannot reject the hypothesis that wage responsiveness to spatial advantages has not increased over the period for unskilled workers. One possibility is that unskilled migration has accelerated over the period due to policy changes discussed above and that this internal labor reallocation has dampened the response of wages for low-skilled workers to enhanced export opportunities.[27]

[26] Estimates from quantile regression models are solutions to the problem of minimizing the sum of absolute residuals. (Koenker and Bassett, 1978; Koenker and Hallock, 2001).

[27] Although a Heckscher–Ohlin framework suggests that capital inflows for export processing in China raise the relative price of skilled workers if such processing is skill intensive, such a result need not hold in other settings. Chakrabarti and Mitra (2010), in a model with asymmetric capital adjustment costs and four goods, show that offshoring-induced capital inflow will not raise the skilled wage premium. Thus, inferences about how deeper economic integration influences domestic skilled and unskilled wages depend on the assumed input–output structure.

**Table 4**
Market access and wages, by education group, 1995 and 2002.

| Dependent variable: log hourly wage | | | | | | |
|---|---|---|---|---|---|---|
| | Skilled | | | Unskilled | | |
| | 1995 | 2002 | Difference | 1995 | 2002 | Difference |
| *Market access* | | | | | | |
| Log MA | 0.270*** | 0.683*** | 0.414*** | 0.313* | 0.653*** | 0.340 |
| | (0.045) | (0.069) | (0.096) | (0.158) | (0.164) | (0.265) |
| *Individual characteristics* | | | | | | |
| Female | −0.074*** | −0.121*** | −0.050* | −0.179*** | −0.263** | −0.084 |
| | (0.017) | (0.020) | (0.027) | (0.039) | (0.092) | (0.075) |
| Years of schooling | 0.022*** | 0.046*** | 0.024** | 0.013 | 0.014 | 0.002 |
| | (0.003) | (0.008) | (0.008) | (0.011) | (0.012) | (0.018) |
| Experience | 0.013*** | 0.004 | −0.008** | 0.023*** | 0.002 | −0.021** |
| | (0.003) | (0.003) | (0.004) | (0.004) | (0.007) | (0.007) |
| Age | 0.064*** | 0.066*** | 0.003 | 0.083** | 0.085** | 0.002 |
| | (0.010) | (0.013) | (0.016) | (0.028) | (0.041) | (0.027) |
| $Age^2$ | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Communist | 0.089*** | 0.084** | −0.005 | 0.016 | 0.089 | 0.073 |
| | (0.024) | (0.032) | (0.041) | (0.068) | (0.054) | (0.082) |
| *Ownership dummies* | | | | | | |
| State-owned | 0.155*** | 0.152*** | −0.003 | 0.167*** | 0.016 | −0.151 |
| | (0.031) | (0.038) | (0.035) | (0.042) | (0.090) | (0.091) |
| Foreign-owned | 0.307*** | 0.341*** | 0.034 | 0.371*** | 0.233 | −0.138 |
| | (0.046) | (0.088) | (0.092) | (0.104) | (0.143) | (0.151) |
| Occupation dummies | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector-province fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | | 7382 | | | 1831 | |
| $R^2$ | | 0.54 | | | 0.47 | |

Notes: Robust standard errors in parentheses; Private-owned is the left out category.
* 10% significance.
** 5% significance.
*** 1% significance.

### 4.4. Real wages and market access

Consideration of internal labor migration leads us to estimate the relationship between spatial advantages and real wages. We define the real wage as the market wage divided by the minimum expense needed to reside in the city where the worker lives. [28] NEG theory predicts that if workers are sufficiently mobile across cities, spatial advantages will be shifted away from workers to the owners of fixed factors, through higher land prices and higher rents for housing and other land improvements.

Table 5a provides the results of regressing market access and individual characteristics on this measure of the real wage, including in the estimating equation provincial fixed effects. The first panel shows coefficient estimates for the full sample. Market access is found to be an insignificant determinant of real wages in both years, with the point estimate actually negative for 1995. The estimated coefficient for 2002 is positive, however, and the difference in the coefficients across the two years is significant. This evidence suggests that intra-provincial migration is sufficiently large in both periods for fixed factors to gain the within-province advantages created by market access. Looking at panel (2) and panel (3) in Table 5a, we see that market access is not a significant determinant of within-province real wages differences for skilled workers or unskilled workers. Mobility of both types of workers appears to be sufficient to shift rents forward to other factor owners.

Table 5b shows coefficient estimates for the same regression, but omitting provincial fixed effects in the estimation. In contrast to the results obtained looking within provinces, we find that workers in provinces with better market access receive higher wages net of living costs than those in less advantageous locations. The first panel provides results for the full sample and we

[28] Average monthly minimum living cost for each city in the sample is calculated as follows. The CHIP surveys report the monthly minimum living expenses for each individual's household. The survey in 1995 asks: "According to actual conditions in your household, please estimate the monthly cost of maintaining a minimum standard of living for the whole family in 1995 (Yuan)". The survey question in 2002 asks: "According to your actual situation, please estimate how much household living expenses per month are needed to maintain a minimum living standard in 2002? (Yuan)". The monthly minimum household living expenses are averaged for every city in each sample to arrive at the average monthly minimum living cost. Nominal monthly living cost in 2002 is converted to 1995 yuan using province level Consumer Price Index (CPI) obtained from various issues of the China Statistical Yearbook. The monthly minimum living cost is converted to hourly terms using the total number of hours worked per month.

**Table 5a**
Market access and wages net of living costs, 1995 and 2002, including provincial fixed effects.

| Dependent variable: (log hourly wage − log hourly living costs)[a] | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | All | | | Skilled | | | Unskilled | | |
| | 1995 | 2002 | Difference | 1995 | 2002 | Difference | 1995 | 2002 | Difference |
| *Market access* | | | | | | | | | |
| Log MA | −0.022 | 0.147 | 0.170** | −0.043 | 0.161 | 0.204** | 0.012 | 0.143 | 0.131 |
| | (0.069) | (0.094) | (0.068) | (0.079) | (0.103) | (0.093) | (0.144) | (0.166) | (0.235) |
| *Individual characteristics* | | | | | | | | | |
| Female | −0.096*** | −0.118*** | −0.021 | −0.076*** | −0.103*** | −0.027 | −0.169*** | −0.233** | −0.064 |
| | (0.021) | (0.020) | (0.026) | (0.019) | (0.022) | (0.031) | (0.043) | (0.097) | (0.080) |
| Years of schooling | 0.014*** | 0.035*** | 0.021*** | 0.013*** | 0.040*** | 0.027*** | 0.009 | 0.007 | −0.001 |
| | (0.004) | (0.006) | (0.005) | (0.003) | (0.008) | (0.008) | (0.012) | (0.017) | (0.023) |
| Experience | 0.015*** | 0.000 | −0.015*** | 0.011*** | 0.001 | −0.010** | 0.023*** | 0.000 | −0.023** |
| | (0.002) | (0.003) | (0.004) | (0.003) | (0.004) | (0.004) | (0.004) | (0.008) | (0.009) |
| Age | 0.067*** | 0.065*** | −0.003 | 0.066*** | 0.065*** | −0.001 | 0.083*** | 0.087** | 0.004 |
| | (0.010) | (0.014) | (0.014) | (0.010) | (0.013) | (0.016) | (0.024) | (0.043) | (0.032) |
| $Age^2$ | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Communist | 0.074*** | 0.088** | 0.013 | 0.095*** | 0.081** | −0.014 | −0.013 | 0.124** | 0.137 |
| | (0.023) | (0.032) | (0.037) | (0.022) | (0.035) | (0.041) | (0.058) | (0.059) | (0.082) |
| *Ownership dummies* | | | | | | | | | |
| State-owned | 0.150*** | 0.143*** | −0.007 | 0.153*** | 0.144*** | −0.009 | 0.142*** | 0.053 | −0.090 |
| | (0.020) | (0.043) | (0.042) | (0.029) | (0.044) | (0.042) | (0.041) | (0.098) | (0.103) |
| Foreign-owned | 0.269*** | 0.316*** | 0.047 | 0.271*** | 0.317*** | 0.046 | 0.280*** | 0.253* | −0.027 |
| | (0.040) | (0.083) | (0.096) | (0.045) | (0.085) | (0.086) | (0.087) | (0.136) | (0.150) |
| Occupation dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector-province fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 9213 | | | | 7382 | | | 1831 | |
| $R^2$ | 0.17 | | | | 0.16 | | | 0.25 | |

Notes: Robust standard errors in parentheses. Private-owned is the left out category.
[a] Hourly living cost is calculated as monthly minimum living cost divided by the number of hours worked per month.
* 10% significance.
** 5% significance.
*** 1% significance.

see that the elasticity of the real wage with respect to market access is 0.059 in 1995 and 0.181 in 2002 and that the difference is significant across the two years. Very similar elasticities are estimated for both the sample of skilled workers and the sample of unskilled workers. While we cannot rule out the influence of unobserved provincial characteristics, these results suggest that barriers to cross-provincial migration are sufficiently high for all workers, skilled and unskilled, to deter the real-wage equalizing impact of internal migration. Workers in locations with better market access earn wages that more than compensate for the higher living costs of spatially advantaged cities. These results are consistent with the view that market access is an independent cause of coastal-inland income inequality.

### *4.5. Heterogeneity by employer type*

We are particularly interested in variation in responsiveness to market access in the wages paid by various types of employers. As noted earlier, between 1995 and 2002, state-owned enterprises (SOEs) shed excess labor and were freed of many administrative controls on wage setting. Hering and Poncet (2010) found a significant relationship between market access and the wages of worker employed by SOEs, but also that the estimated elasticity was substantially smaller than that found for private and foreign enterprises. Table 6 shows our estimates of market access elasticities by firm ownership type for 1995 and 2002. Our results for 1995 are qualitatively similar to those obtained by Hering and Poncet – market access is a significant determinant of wages for all firm types, but the responsiveness of wages in SOEs is significantly less than that of private and foreign enterprise.

For all ownership types, the estimated elasticity of the wage with respect to market access was larger in the 2002 sample than in the 1995 sample and the difference across years was statistically significant for both SOEs and private firms. While the estimated coefficient for foreign firms is larger for the 2002 sample than for the 1995 sample, this difference is not statistically significant. Nevertheless, in 2002 wages paid by foreign enterprises are the most responsive to cross-city differences in market access.

As expected given the changes in wage-setting controls, the 2002 sample indicates a much stronger response of SOE wages to market access than does the 1995 sample. The estimated elasticity more than doubles between the two years. This difference is highly significant and is consistent with increasingly market-determined wage levels within the state sector. By 2002, the estimated elasticity for SOEs is two-thirds that estimated for private and foreign firms, compared to an estimated elasticity that is half as large as its private counterparts in 1995.

**Table 5b**
Market access and wages net of living costs, 1995 and 2002, excluding provincial fixed effects.

| Dependent variable: (log hourly wage − log hourly living costs)[a] | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | All | | | Skilled | | | Unskilled | | |
| | 1995 | 2002 | Difference | 1995 | 2002 | Difference | 1995 | 2002 | Difference |
| *Market access* | | | | | | | | | |
| Log MA | 0.059*** | 0.181*** | 0.122*** | 0.072*** | 0.188*** | 0.116*** | 0.040 | 0.161*** | 0.121*** |
| | (0.019) | (0.043) | (0.025) | (0.010) | (0.043) | (0.033) | (0.051) | (0.025) | (0.030)* |
| *Individual characteristics* | | | | | | | | | |
| Female | −0.095*** | −0.110*** | −0.015 | −0.073*** | −0.096*** | −0.023 | −0.164*** | −0.181*** | −0.016 |
| | (0.016) | (0.019) | (0.029) | (0.015) | (0.017) | (0.030) | (0.015) | (0.046) | (0.049) |
| Years of schooling | 0.015 | 0.035*** | 0.019*** | 0.016*** | 0.043*** | 0.027** | 0.021*** | 0.004 | −0.018 |
| | (0.029) | (0.004) | (0.002) | (0.002) | (0.011) | (0.011) | (0.003) | (0.008) | (0.011) |
| Experience | 0.014*** | 0.002 | −0.013** | 0.010*** | 0.003 | −0.007 | 0.021*** | 0.001 | −0.019** |
| | (0.001) | (0.005) | (0.004) | (0.001) | (0.004) | (0.004) | (0.003) | (0.010) | (0.007) |
| Age | 0.065*** | 0.070*** | 0.005 | 0.066*** | 0.071*** | 0.005 | 0.072*** | 0.085*** | 0.013 |
| | (0.006) | (0.015) | (0.010) | (0.011) | (0.017) | (0.006) | (0.009) | (0.019) | (0.010) |
| Age$^2$ | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Communist | 0.078*** | 0.089** | 0.011 | 0.102*** | 0.078** | −0.024 | −0.026 | 0.141*** | 0.167*** |
| | (0.015) | (0.036) | (0.039) | (0.020) | (0.034) | (0.047) | (0.028) | (0.030) | (0.012) |
| *Ownership dummies* | | | | | | | | | |
| State-owned | 0.175*** | 0.136*** | −0.039 | 0.177*** | 0.143*** | −0.034 | 0.172*** | 0.085*** | −0.087** |
| | (0.014) | (0.020) | (0.033) | (0.016) | (0.023) | (0.033) | (0.024) | (0.012) | (0.032) |
| Foreign-owned | 0.245*** | 0.315*** | 0.069 | 0.256*** | 0.319*** | 0.063 | 0.195*** | 0.238*** | 0.043 |
| | (0.014) | (0.036) | (0.049) | (0.015) | (0.023) | (0.037) | (0.054) | (0.014) | (0.068) |
| Occupation dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 9213 | | | | 7382 | | | 1831 | |
| $R^2$ | 0.59 | | | | 0.60 | | | 0.59 | |

Notes: Robust standard errors in parentheses. Private-owned is the left out category.
[a] Hourly living cost is calculated as monthly minimum living cost divided by the number of hours worked per month.
* 10% significance.
** 5% significance.
*** 1% significance.

### 4.6. Controlling for agglomeration

NEG theory predicts that freer trade will induce greater concentration of production as firms locate in places with the best access to foreign markets. Thus, increased agglomeration of economic activities may be seen as a by-product of the same forces that relate local wages to foreign market access. However, urban economic theory provides other reasons for firms to cluster spatially and these forces also raise local wages. For example, China's progressive opening to foreign trade led to significant clustering of foreign firms in eastern cities, and these cities are closer to each other than to inland cities. External economies from these foreign firms may have driven wages higher in eastern provinces, even if there is no response of firms to differences across Chinese cities

**Table 6**
Estimated market access elasticity, by firm ownership type, 1995 and 2002.

| Dependent variable: log hourly wage | | | |
|---|---|---|---|
| | 1995 | 2002 | Difference |
| *Market access* | | | |
| Log MA × state-owned | 0.215*** | 0.463*** | 0.248*** |
| | (0.053) | (0.070) | (0.074)* |
| Log MA × private-owned | 0.409*** | 0.610*** | 0.201** |
| | (0.043) | (0.095) | (0.090) |
| Log MA × foreign-owned | 0.423*** | 0.624*** | 0.200 |
| | (0.080) | (0.143) | (0.173) |
| Observations | | 9213 | |
| $R^2$ | | 0.53 | |

Notes: Robust standard errors in parentheses. Regressions include individual characteristics, ownership, occupation, and province-sector dummies.
* 10% significance.
** 5% significance.
*** 1% significance.

in access to markets. If such a mechanism is at work, we may inappropriately attribute the effect of external economies to market access.

To the extent that these agglomerative factors occur at the provincial and sector level, we have controlled for them by our inclusion of province and sector fixed effects. These fixed effects, however, cannot capture the wage effect of spillovers created by the density of economic activity at the city level. Therefore, we include in our wage regressions a number of alternative measures of city-level economic activity. The first measure is employment density, which we create using firm level data from China's Annual Survey of Industrial Production (ASIP). It is defined as total employment in the city per square kilometer. The second measure is firm density, also created using ASIP data, and defined as the number of industrial firms located in the city per square kilometer. Our last measure is city population, a measure of city size.

Results of our wage regressions including these agglomeration measures are shown in Table 7. Panel (1) provides estimated coefficients when we include log employment density, panel (2) provides results when we include firm density, and panel (3) when we include population. In all cases, market access remains a significant determinant of urban wages in both sample years. Interestingly, the agglomeration measures are negatively and insignificantly related to wages in 1995. By 2002, however, a positive relationship between economic density and wages appears and this relationship is significant for employment density, as seen in panel (1). The estimated elasticity of wages with respect to employment density and firm density become significantly larger between 1995 and 2002 and inclusion of either measure reduces the estimated coefficient for market access in 2002. As noted above, a city with market access advantages may be predicted to experience both rising wages and greater firm or employment density in response to liberalization, so these results could indicate this process at work in China. Nevertheless, we cannot rule out the possibility that agglomerative forces other than market access are driving the increased spatial component of urban wage inequality, even as market access remains a significant determinant of wages. Finally, we note that the impact of market access on wages is unchanged by inclusion of a control for

**Table 7**
Market access and wages, controlling for agglomeration, 1995 and 2002.

| Dependent variable: log hourly wage | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | | | (2) | | | (3) | | |
| | 1995 | 2002 | Difference | 1995 | 2002 | Difference | 1995 | 2002 | Difference |
| *Market access* | | | | | | | | | |
| Log MA | 0.396*** | 0.422*** | 0.025 | 0.525** | 0.446*** | −0.080 | 0.288*** | 0.651*** | 0.362*** |
| | (0.103) | (0.119) | (0.175) | (0.184) | (0.130) | (0.239) | (0.046) | (0.057) | (0.060) |
| *Agglomeration variables* | | | | | | | | | |
| Log employment density | −0.031 | 0.066* | 0.098* | – | – | – | – | – | – |
| | (0.027) | (0.036) | (0.053) | | | | | | |
| Log firm density | – | – | – | −0.070 | 0.061 | 0.130* | – | – | – |
| | | | | (0.048) | (0.039) | (0.070) | | | |
| Log population | – | – | – | – | – | – | 0.011 | 0.045 | 0.034 |
| | | | | | | | (0.059) | (0.072) | (0.082) |
| *Individual characteristics* | | | | | | | | | |
| Female | −0.098*** | −0.137*** | −0.039 | −0.100*** | −0.137*** | −0.038 | −0.099*** | −0.138*** | −0.039 |
| | (0.018) | (0.020) | (0.024) | (0.018) | (0.020) | (0.024) | (0.018) | (0.020) | (0.024) |
| Years of schooling | 0.022*** | 0.041*** | 0.019*** | 0.023*** | 0.041*** | 0.019*** | 0.022*** | 0.042*** | 0.020*** |
| | (0.004) | (0.005) | (0.006) | (0.004) | (0.005) | (0.005) | (0.003) | (0.005) | (0.006) |
| Experience | 0.016*** | 0.003 | −0.013*** | 0.017*** | 0.003 | −0.013*** | 0.017*** | 0.003 | −0.013*** |
| | (0.002) | (0.003) | (0.004) | (0.002) | (0.003) | (0.004) | (0.002) | (0.003) | (0.004) |
| Age | 0.068*** | 0.068*** | 0.000 | 0.067*** | 0.068*** | 0.000 | 0.067*** | 0.068*** | 0.001 |
| | (0.011) | (0.013) | (0.014) | (0.011) | (0.013) | (0.015) | (0.011) | (0.013) | (0.015) |
| $Age^2$ | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 | −0.001*** | −0.001*** | 0.000 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Communist | 0.076*** | 0.087** | 0.011 | 0.074*** | 0.087*** | 0.014 | 0.075*** | 0.086*** | 0.011 |
| | (0.025) | (0.029) | (0.036) | (0.024) | (0.029) | (0.036) | (0.024) | (0.029) | (0.036) |
| *Ownership dummies* | | | | | | | | | |
| State-owned | 0.154*** | 0.137*** | −0.017 | 0.155*** | 0.140*** | −0.015 | 0.158*** | 0.140*** | −0.018 |
| | (0.024) | (0.037) | (0.035) | (0.023) | (0.036) | (0.034) | (0.024) | (0.037) | (0.033) |
| Foreign-owned | 0.318*** | 0.323*** | 0.005 | 0.316*** | 0.326*** | 0.010 | 0.316*** | 0.327*** | 0.011 |
| | (0.036) | (0.083) | (0.010) | (0.039) | (0.084) | (0.102) | (0.037) | (0.081) | (0.099) |
| Occupation dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sector-province fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 9078 | | | | 9213 | | | 9213 | |
| $R^2$ | 0.53 | | | | 0.53 | | | 0.53 | |

Notes: Robust standard errors in parentheses. Private-owned is the omitted category.
* 10% significance.
** 5% significance.
*** 1% significance.

city size. Panel 3 shows that the estimated market access elasticity remains essentially unchanged from that estimated without including population and the increase in the elasticity across the two samples is statistically significant.

### 4.7. Decomposing the change in mean wages

A decomposition of the change in the mean wage indicates that location is an economically important factor explaining the growth in average wages between 1995 and 2002. The Mincer wage equations take the form $y = f(x)$ where $y$ is the logarithm of individual hourly wages in 1995 yuan and $x$ is the vector of explanatory variables. Using the wage equations, displayed in Table 2, the growth of wages between 1995 and 2002 can be decomposed as,

$$\bar{y}_{2002} - \bar{y}_{1995} = f_{2002}(\bar{x}_{2002} - \bar{x}_{1995}) + (f_{2002} - f_{1995})\bar{x}_{1995},$$

where the first term reflects the effect of differences in the characteristics and the second the effect of differences in the coefficients between 2002 and 1995. We also carry out the alternative decomposition as,

$$\bar{y}_{2002} - \bar{y}_{1995} = f_{1995}(\bar{x}_{2002} - \bar{x}_{1995}) + (f_{2002} - f_{1995})\bar{x}_{2002},$$

Results of this Oaxaca–Blinder decomposition are shown in Table 8.

This decomposition allows us to measure the percentage change in mean wages that is due to the growth in the size of distance-weighted markets access (mean characteristics) and the percentage due to increased importance of market access as a determinant of wages (coefficient estimates). Looking first at the top half of Table 8, we see that between 7.5% (using 2000 as the base year) and 17.5% (using 1995 as the base year) of the difference in mean wages can be explained by the growth in the size of potential markets. As seen in Table 1, the size of markets accessible to workers in the sample grew by 76% between 1995 and 2002. By either decomposition method, changes in accessible market size are the largest determinant of the change in mean wages attributable to changes in worker characteristics.

Mean wages also changed because of differences across the two years in the market value of given worker characteristics. Looking at the bottom half of Table 8, we see that between 22.7% (using 1995 as the base year) and 12.8% (using 2002 as the base year) of the change in mean wages is due to the heightened responsiveness of wages to market access. The increased importance of market access explains as much of the growth in mean wages as does the increased importance of education. Thus, access to richer markets and heightened importance of that market access for factor prices are important explanations for the growth in average urban wages over the period.

**Table 8**
Blinder–Oaxaca decomposition of the increase in mean wages, 1995–2002.

| Percentage due to: | (1) | (2) |
|---|---|---|
| *Mean characteristics* | | |
| Total | 10.67 | 25.93 |
| Market access | 7.51 | 17.52 |
| Female | 0.43 | 0.60 |
| Communist party membership | 0.29 | 0.34 |
| Education | 0.99 | 1.89 |
| Experience | 2.51 | 0.53 |
| Occupation | 0.01 | 0.32 |
| Age | 0.19 | 2.16 |
| Ownership | −2.16 | −1.84 |
| Sector-province | 0.90 | 4.41 |
| *Coefficients* | | |
| Total | 89.33 | 74.07 |
| Market access | 22.76 | 12.75 |
| Female | −1.50 | −1.67 |
| Communist party membership | 0.23 | 0.19 |
| Education | 19.17 | 18.27 |
| Experience | −24.59 | −22.61 |
| Occupation | −19.68 | −19.99 |
| Age | 19.04 | 17.07 |
| Ownership | −0.79 | −1.11 |
| Sector-province | −7.09 | −10.60 |

Notes: The explained variable is the difference in mean wages, expressed in 1995 yuan, between 1995 and 2002; Decompositions based on results from Table 2, column 3; column 1 shows the differences in the predictors between 1995 and 2002 weighted by the coefficients and predictor levels in 1995, and column 2 shows the differences in the predictors between 1995 and 2002 weighted by the coefficients and predictor levels in 2002.

## 5. Conclusion

Motivated by the far-reaching liberalization of trade and labor markets undertaken by China after 1995, we measure the extent to which the influence of market access on wages strengthened and influenced wage growth over the subsequent period. We find that the strength of this city characteristic on urban wages did increase and that it contributed in an economically significant way to the growth in average wages between 1995 and 2002. Our estimates indicate that market access influenced wages of both skilled and unskilled workers, with the estimated elasticity almost identical for the two groups. Therefore, while a stronger market access effect is associated with higher average wages, this spatial factor does not seem to be an explanation for a rising skill premium within cities.

Our results for differences in wages adjusted for cost of living indicate the need for further research on how migration influences spatial wage differences. We find no empirical support for the existence of within-province differences in net-of-living-cost wages caused by market access. Thus, our results are suggestive of sufficient within-province labor migration and sufficiently low segmentation between markets for resident and migrant workers to eliminate difference across cities in the same province. When we omit provincial fixed effects from the Mincerian wage regressions, however, our estimates open the possibility of other inferences. Coefficient estimates suggest that barriers to inter-provincial migration are large enough to prevent cross-province labor flows from eliminating the advantages to workers of coastal locations. These findings point to the relevance of further study of cross-province income differences and shifts between owners of mobile and fixed factors.

We also find that while the wages of workers employed by state-owned enterprises had the lowest responsiveness to intercity differences in market access, it is this group of workers for whom the market access effect increased the most. The evidence is consistent with other measures of increasingly market-driven behavior by state enterprises. Overall, our study provides evidence directly from the wages of urban workers that the structure of Chinese product and labor markets has been significantly influenced by ongoing commercial and regulatory policy changes.

## Data Appendix

| Province | Cities in 1995 CHIP | Cities in 2002 CHIP |
|---|---|---|
| Beijing | Beijing | Beijing |
| Shanxi | Changzhi, Datong, Fenyang, Taiyuan, Yangquan | Changzhi, Datong, Fenyang, Hunyuan, Taiyuan, Yuncheng, Xing |
| Liaoning | Dalian, Jinzhou, Shenyang | Changtu, Dalian, Jinzhou, Shenyang, Wafangdian |
| Jiangsu | Changzhou, Dafeng, Nanjing, Nantong, Suqian, Taixing, Wuxi, Xuzhou, Yixing | Dafeng, Nanjing, Nantong, Taixing, Suqian, Wuxi, Xuzhou, Yangzhou, Yixing |
| Anhui | Bengbu, Bozhou, Hefei, Huainan, Tongcheng, Wuhu | Bengbu, Bozhou, Hefei, Huainan, She, Wuhu |
| Henan | Huixian, Kaifeng, Pingdingshan, Xiangcheng, Xinxiang, Zhengzhou | Gushi, Hua, Huixian, Kaifeng, Pingdingshan, Xiangcheng, Xinxiang, Zhengzhou |
| Hubei | Honghu, Huangshi, Macheng, Tianmen, Wuhan, Xiangfan | Honghu, Jingzhou, Wuhan, Xiangfan, Xianning, Xishui, Yichang |
| Guangdong | Foshan, Guangzhou, Huizhou, Puning, Shenzhen, Shunde, Zhanjiang, Zhaoqing | Foshan, Guangzhou, Huizhou, Puning, Shaoguan, Shunde, Zhanjiang, Zhaoqing |
| Chongqing | | Chongqing |
| Sichuan | Chengdu, Guangyuan, Leshan, Luzhou, Zigong | Chengdu, Emeishan, Guanguyuan, Luzhou, Nanchong, Neijiang |
| Yunnan | Dali, Geiju, Kunming, Xuanwei | Dali, Geiju, Kunming, Lijiang Naxi, Pu'er Hani-Yi, Xuanwei |
| Gansu | Lanzhou, Pingliang, Wuwei | Lanzhou, Pingliang, Wuwei |

Note: These are cities in the final samples.

*Sectors included in 1995*: (i) industry; (ii) transportation, communications, posts and telecommunications; and (iii) commerce and trade, restaurants and catering, materials supply, marketing and warehousing.

*Sectors included in 2002*: (i) manufacturing; (ii) transportation, storage, post office and communication; and (iii) wholesale, retail and food services.

*Ownership categories in 1995*: state-owned enterprises at the central or provincial level (central SOEs), local publicly owned enterprises (local SOEs), urban collective enterprises, private enterprises, foreign-invested enterprises, Sino-foreign joint ventures, and others which includes township and village enterprises.

*Ownership categories in 2002*: state-owned enterprises at the central or provincial level (central SOEs), local publicly owned enterprises (local SOEs), urban collective enterprises, private enterprises, foreign-invested enterprises, Sino-foreign joint ventures, and others which includes rural private and rural individual enterprises and other share-holding companies. An additional category, state share-holding company, appears in the 2002 individual survey. We experimented including this category with central SOEs and local SOEs.[29] Our benchmark results are invariant to the alternative categorizations. Therefore, we chose to display results which include this category with central SOEs to minimize the number of tables.

[29] Huang (2008) provides evidence that share-holding companies are mostly privatized state-owned firms at local levels of government. Levels of local government in China include the province, prefecture, county, township and village.

We collapse the various ownership categories in each of the survey years into three broad types — state-owned, private-owned, and foreign-owned enterprises. State-owned includes central and local SOEs; private-owned includes urban collective enterprises, private enterprises, and other enterprises; and foreign-owned includes foreign-invested enterprises, and Sino-foreign joint ventures.

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