

## Urban Unemployment Duration Analysis in Post-Reform China

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### ABSTRACT

We link China's accession to the WTO with the unemployment duration of urban jobseekers. Using data from the Rural Urban Migration in China (RUMiC) of a survey conducted in early 2008, we construct an inflow sample of those who started a period of joblessness between January 2002 and December 2007 to estimate the unemployment duration for local urban jobseekers. Deploying the Cox Proportional Hazards and Accelerated Failure Time models, we find reduced trade policy uncertainty significantly shortens unemployment duration, with hazard-rate varying across internal in-migrants and urban natives, due to variations in job accessibility and reservation wages.

### KEYWORDS

Unemployment Duration, Trade Policy Uncertainty, Hukou, Survival Analysis

### INTRODUCTION

As the fastest growing developing economy over the last 30 years, China has attracted the attention of academia. China's remarkable story of structural change is widely attributed to its integration into the global economy, propelled by its 2001 accession to the World Trade Organization (WTO) (Feenstra and Wei, 2010; Autor, et al., 2013; Handley and Limao, 2017; Pierce and Schott, 2016; Facchini, et al., 2019). Whilst the consequent ease of access of Chinese exports to world markets has encouraged demand for cheap local labour within Chinese labour markets, structural differences between different groups of jobseekers across various dimensions (i.e., relative income, training and education—with or without urban household registration known as “Hukou”<sup>1</sup>—social networks, family and social relationships) have exaggerated the conflicts among displaced jobseekers. One pertinent aspect here is unemployment durations across local urban jobseekers with Hukou registration vis-à-vis internal in-migrants without it, owing to China's attainment of the Permanent

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<sup>1</sup> Hukou is a residency permit introduced by the Chinese government. It predicts not only the location of registration and work of everyone residing in China but also access to local public services. Since the structural liberalization policies of the 1980s, people can reside and work outside of their hukou registration place, but the acquisition of local hukou registration and attainment of urban rights and jobs by the internal in-migrants aiming to be located in the urban area are still denied since local governments prioritize the needs of registered urban residents over in-migrants. See Zhang (2010).

Normal Trade Relations (PNTR) status.<sup>2</sup> It is vital to understand the impact of reduced trade policy uncertainty and consequent changes in trade patterns with the United States (US) on China's local labour market outcomes regarding unemployment duration across internal in-migrants and local natives of the urban Chinese population, both because of the scope of the "Great Migration" within China (Scheineson, 2009) and because of the massive impact of liberal trade policy regime in the US and the "Great Migration" on the Chinese economy and society (Facchini, et al., 2019). Focusing on the favourable trade policy shock as a natural experiment and by deploying suitable survival models, this paper exemplifies the causal impact of this policy change on unemployment duration through the changes in determinants of duration.

Indeed, China's attainment of PNTR status could have resulted in the continuation of Chinese exports to the US at the MFN (Most Favoured Nation) rate, leading to a reduction of uncertainty associated with the tax rate that these Chinese exporters face in the US import markets.<sup>3</sup> This may have affected unemployment durations across these two groups, primarily through the following channels: On one hand, reduction in trade policy uncertainty might have incentivized the local Chinese exporters to boost production to serve the US market, resulting in a greater demand for cheap labour. This could contribute to a reduction in the unemployment duration due to new job creation and improved accessibility in the export industries. On the other hand, the differences in reservation wages between the two groups, viz., the local urban jobseekers and the internal in-migrants and the prevalence of differential access of migrants (vis-à-vis the local urban jobseekers) to public goods and services might have impacted differently the chances of getting a new job among these two groups. Therefore, the question remains about what the implication of China's attainment of PNTR status should be, in terms of both magnitude and direction, on the unemployment duration among these two groups of displaced local jobseekers, viz., the internal in-migrants versus local urban jobseekers in China.<sup>4</sup>

To examine this, we first measure the reduction in trade policy uncertainty in the Chinese exportable industries selling to the US<sup>5</sup> (following Pierce and Schott, 2016; see section 3 for the

<sup>2</sup> The US Congress approved a bill granting PNTR status to China in October 2000. PNTR came into effect upon China's WTO accession in December 2001 and was implemented on 1 January 2002. PNTR was different from the traditional trade facilitation measures in the sense that it eliminated a major source of uncertainty in US-China trade relations instead of changing the actual USA tariff rates applied to Chinese goods. It is not that before PNTR, US imports from China were subject to higher import tariff rates than were available to most other US trading partners affiliated to the WTO. However, since the US government classified China as a non-market economy, continued admittance to those low rates necessitated annual re-approval by the president, which could always be overturned by Congress. Approval of these renewals on the annual basis became more unreliable after the Tiananmen Square incident in 1989 and remained so throughout periods of tension between the US and China in the 1990s. Notably, if any renewal were to fail, US tariffs on most Chinese imports would have increased substantially, from 4% to 37% for the average manufacturing industry. The change in China's PNTR status ended the uncertainty associated with annual renewals of China's normal-trade-relation status with the US, abolishing any option value of waiting for the US and any potential sunk costs that Chinese exporting firms used to incur for exporting to US markets. This led to a substantial reduction in expected US import tariffs on Chinese goods. See Pierce and Schott (2019).

<sup>3</sup> Hereafter, we shall refer this as a 'reduction in trade policy uncertainty'.

<sup>4</sup> While the other policy changes, such as the change in the *Hukou* system over the years, institutional changes (for example, change in the one-child policy) or the effects of an ageing population (that could affect migration behaviour) are important, but none of these alternative policy changes is specific to the urban population. Therefore, any of these policy changes cannot act as a possible exogenous determinant of unemployment durations among the urban population. See Facchini, et al. (2019).

<sup>5</sup> Given the sheer importance of the US market as primary destination of the Chinese exports (Facchini, et al., 2019), we are using China's attainment of the PNTR status and the consequent liberal trade policies faced in the US by the Chinese exporters as the exogenous policy that triggered the local labour demand shock and unemployment duration.

details); then, following Facchini, et al. (2019),<sup>6</sup> we utilize this measure of reduced trade policy uncertainty (aggregated at the prefecture-levels) to empirically predict unemployment duration in the Chinese labour market. Using the data on Rural Urban Migration in China (RUMiC) from an early 2008 survey, we construct an inflow sample of the unemployed made up of those who began a period joblessness between January 2002 and December 2007. This allowed us to empirically estimate unemployment duration for local jobseekers in urban China during the post-WTO regime. Utilising a suite of models under “*survival analysis*” (viz., Proportional Hazard (PH) and Accelerated Failure Time (AFT) models, Addison and Portugal, 1989; Van den Berg, 2001; Kupets, 2006; Liu and Yu, 2019), we show how the positive boost in exports following reduced trade policy uncertainty would affect China’s domestic labour market by altering the reallocation of labour, income and employment status for the internal in-migrants and the urban natives within China’s internal labour market and thereby contributing to the conflict.

Our research makes a dual contribution to the existing literature. Initially, we delineate between two distinct groups of jobseekers in China: those who relocate within urban areas without obtaining *Hukou* (registration), termed as “non-*Hukou*” internal “in-migrants,” and those who have already acquired urban *Hukou*, termed as urban *Hukou* “natives.” Subsequently, we delve into two main aspects. Firstly, we explore the impact of reducing trade policy uncertainty on internal migration to urban prefectures, examining the differential responses of ‘in-migrants’ versus ‘native’ jobseekers. Secondly, we investigate how structural disparities such as gender, education, health status, training, experience, household income and expenditure, and the size of social networks influence internal migration patterns. Due to the limitations of traditional regression techniques in handling censored duration data (Heckman and Singer, 1982), we employ survival analysis methodology to analyse unemployment duration data in this study. This methodology considers both censored and uncensored observations and accommodates the stochastic nature of the point of censoring. Additionally, we deploy AFT models, with the family of distributions, which has been described in detail by Kalbfleisch and Prentice (1981) and applied by Lancaster (1979), Heckman and Borjas (1980) and Nickell (1979).

The remainder of the paper is organised as follows. Section 2 provides a literature review on the correlation between trade policy adjustments and labour market dynamics. Section 3 presents data and variable descriptions, identification strategy, and transmission mechanisms. Section 4 outlines the econometric/statistical framework applied in the analysis, while Section 5 presents the findings. Lastly, Section 6 concludes with potential avenues for future research.

## LITERATURE REVIEW

The theme of this paper is guided by recent literature that investigates the link between trade reform policies and factor market adjustments in China. In this context, Potlogea and Cheng (2017) utilised cross-city variation in terms of the city-level exposure to foreign markets, owing to preliminary differences in sectoral specialisation and exogenous cross-industry differences in the US tariff cuts for Chinese imports, and owing to the latter’s attainment of PNTR status. The authors then investigated the local development impact on multiple outcome variables, namely changes in urbanisation, output, employment, investment and FDI flows. Recently, Erten and Leight (2021) utilised a panel of Chinese country-specific data to investigate the impact of reduced trade policy uncertainty on China’s structural transformation at the local level. The paper finds that counties exposed to reduced

<sup>6</sup> Facchini, et al. (2019) largely followed Bartik (1991) for determining the local exposure to uncertainty at the level of a Chinese prefecture. It is well-known that the Bartik (1991) method relies upon the strict exogeneity of the industry shares, but as demonstrated by Goldsmith-Pinkham, et al. (2020), this assumption is about exogeneity conditional on observables: viz., the changes in outcome variable(s) rather than in levels of the outcome variable.

uncertainty surrounding their export prospects have experienced substantial expansion of, and increased employment in, the manufacturing sector but reduced employment in the agriculture sector and greater influx of internal in-migration (proxied by an increase in population stock at the county-level) following China's WTO accession.

In a contrasting study, Zi (2020) examined the impact of Chinese liberalisation of trade in intermediate products on the relocation of economic activity across the country and how that contributes to changes in the restrictiveness of the *Hukou* system across provinces. The author then traces out the welfare gains owing to the elimination of the *Hukou* system. In a similar vein, Dai, et al. (2021) recently examined effects of the opening-up of the Chinese economy following the WTO accession on urban household behaviour in China; they found that regions exposed to greater tariff cuts experienced lower wages.

Motivated by the fact that the Chinese economy experienced a reduction in trade policy uncertainty faced by the Chinese exporters in the US import markets and following China's attainment of PNTR status granted by the US, Facchini, et al. (2019) examined the effects of reduction in trade policy uncertainty on internal migration in China. Interestingly, they found that the reduction in trade policy uncertainty is the significant driver behind increasing internal skills for prime working age (16–65 years) migrants (*non-Hukou*) in China.

In this paper, our focus is to explore the impact of exogenous trade policy changes on the unemployment duration across local urban “natives” and in-migrants. In other words, we are concerned about the following two questions:

1. How will the impact of exogenous trade policy changes reduce or accelerate the time required for an unemployed urban jobseeker to find a new job in the presence of relevant independent structural factors, e.g., age, gender, education?
2. Does the chance of finding a new job (viz., the hazard rate) vary with trade policy changes across urban natives and in-migrants owing to differences in independent structural factors?

Survival models have been used extensively in business research, with the aim of studying, for instance, the likelihood of order execution within a limit-order book (LOB) and the factors influencing such likelihood (Pan & Mishra, 2022). A common feature of such data sets is the presence of censoring. Note that these censored observations reveal only partial information and as a result, the traditional regression models may not be suitable in handling them (Heckman and Singer, 1982). A common approach to deal with such censored variables is to use the PH model of Cox (1972). In the context of modelling unemployment duration, however, the PH model has been not so frequent, with a few exceptions (Kupets, 2006).

Most importantly, the underlying assumption of proportional hazards may be violated in many applications. In such cases, a possible option is to extend the traditional PH models (Follmann, et al., 1990). A common alternative in such cases is to use the log-location-scale regression models, which are traditionally known as AFT models in the literature (Kalbfleisch and Prentice, 1981). The AFT models are essentially parametric models that fully specify the underlying distribution of time to an event. Such AFT models have also been used in modelling the unemployment duration (Addison and Portugal, 1989; Bergström and Edin, 1992). Furthermore, these AFT models have been used in survival analysis in other contexts (Pan & Mishra, 2022).

Following this methodological trend, in addition to using the most popular model, viz., the Cox PH models (Kupets, 2006; Carroll, 2006; Zhang, 2010), we have also utilised the AFT class of models under a family of distributions. Our paper explores the relative importance of changing trade policy induced labour relocations, *Hukou* regulations and a host of other relevant factors on unemployment duration

across local in-migrants vis-à-vis urban natives in China, considering differences in gender, education, health status, training, experience, household income and expenditure and relative size of the social network with which the respective household is associated.

## **DATA, IDENTIFICATION STRATEGY, VARIABLES & DESCRIPTIVES, TRANSMISSION MECHANISMS**

### **DESCRIPTIONS OF SURVEY AND DATA**

We use the survey-based data on Rural Urban Migration in China (RUMiC) to execute the empirical illustration.<sup>7</sup> Its large-scale, in-depth topics and longitudinal aspect make RUMiC a unique tool to explore migration and the labour market in China (Akgüç, et al., 2014). The full RUMiC survey is comprised of three parts: Urban Household Survey, Migrant Household Survey and Rural Household Survey. Each survey involves around 5000 households. The Urban Household Survey focuses on urban residents, while the Migrant Household Survey focuses on migrants who are registered in another place but are living in the urban areas at the time of the survey.

Both surveys were conducted in early 2008. Each of these two surveys includes comprehensive information on household and personal characteristics, detailed health status, employment, income, training and education of adults and children, social networks, family and social relationships and life events.

This study utilises the Urban Household Survey and the Migrant Household Survey. We utilise an inflow sample of unemployed comprising those who began an unemployment period between January 2002 and December 2007 and who provided a complete response to the questions about their job search period. Given the unavailability of information about the unemployment periods starting after December 2007, we do not include the jobseekers that started an unemployment period after December 2007 in our sample. Samples are restricted to people aged between 16 and 45 at the time of the survey, considering that most migrant jobseekers are aged under 45.<sup>8</sup> As a result, the total sample size adopted in our study is 4,089, of which 60% are migrant jobseekers.

### **IDENTIFICATION STRATEGY**

To investigate the individual-specific heterogeneity in unemployment durations, we construct the following variables from our sample. First, we consider individual characteristics such as age, age-squared, gender, education, health status, training experience and migration status. Previous studies indicated that younger jobseekers and males were more likely to find jobs compared to older jobseekers and females, and jobseekers with more human capital have a higher re-employment probability than unskilled ones (Foley, 1997; Giles, et al., 2006; Kupets, 2006; Du and Dong, 2008). Second, we consider household characteristics such as the size of social network and household income/expenditures. Households with large social networks make it easier for unemployed members to find a job, while a wealthier household sets a higher reservation wage for the unemployed member of the family, in turn leading to unemployment duration persistence (Giles, et al., 2006; Kupets, 2006; Zhang, 2010). In addition to these individual and household characteristics, we use other appropriate measures (as explained in Table 1) of:

<sup>7</sup> The survey was jointly initiated by a group of researchers from the Australian National University, the University of Queensland, and the Beijing Normal University and was supported by the Institute for the Study of Labor (IZA), which provides Scientific Use Files. More details on the RUMiC survey can be found at <https://datasets.iza.org/dataset/58/longitudinal-survey-on-rural-urban-migration-in-china>.

<sup>8</sup> Thanks to the *Hukou* System in China, most migrant jobseekers only stay temporarily in the receiving regions and will return to the sending regions later. According to our data, the mean (median) value of migration duration is about 7 (6) years.

1. the differences in local labour demand conditions (captured by unemployment rate); and
2. the reduced policy uncertainty surrounding the Chinese export markets after China's accession to the WTO (Pierce and Schott, 2016; Facchini, et al., 2019).

*The causal identification channel* linking the reduction in trade policy uncertainty on exports from China to the United States to the heterogeneity in unemployment durations can be elucidated as follows:

The reduction in trade policy uncertainty, consequent to China's attainment of PNTR status with the United States, engenders differential impacts on the local labour market dynamics, thereby influencing unemployment durations. Specifically, we observe that a diminished level of uncertainty in trade policies amplifies labour demand shocks within the local labour market. This heightened labour demand is found to correlate with shorter durations of unemployment among affected individuals.

Furthermore, our analysis reveals a notable discrepancy in unemployment durations between migrant workers and urban natives. This disparity is multifaceted and can be attributed to two distinct mechanisms.

Firstly, the *Hukou* System in China confers certain privileges to urban natives in accessing local labour markets, thereby placing migrant workers at a disadvantage. Consequently, migrant workers are likely to possess lower reservation wages relative to their urban native counterparts, leading to shorter durations of unemployment. This causal pathway underscores the role of differential reservation wages influenced by systemic biases in the urban labour market.

Alternatively, an alternative explanation posits that migrant workers exhibit greater job mobility across sectors and are more attuned to job opportunities compared to urban natives. In this scenario, the lower mobility costs and heightened sensitivity to job openings among migrant workers facilitate quicker re-employment, thereby contributing to shorter unemployment durations. This rationale underscores the significance of enhanced job accessibility as a driving force behind the observed variation in unemployment durations between migrant workers and urban natives amidst the reduction in trade policy uncertainty.

#### OTHER FACTORS BEHIND HETEROGENEOUS IMPACT OF REDUCED TRADE UNCERTAINTY ON UNEMPLOYMENT DURATIONS

- Skill-differentiation and training: Disparities in skill levels and access to training opportunities between urban natives and migrant workers further contribute to differential outcomes in unemployment durations.
- Age differentiations: Variations in age demographics between the two groups may also influence their respective labour market experiences and unemployment durations.
- Size of social network: The presence of large social networks within households facilitates quicker re-employment for unemployed members.
- Household income/expenditures: Wealthier households may sustain longer unemployment durations due to their ability to set higher reservation wages for the unemployed member(s) of their respective families, in turn leading to unemployment duration persistence.

In order to quantitatively illustrate this causal identification channel, we start with elucidating the methodological intricacies underlying our quantification of the abovementioned reduction in trade policy uncertainty encountered by Chinese exporters targeting the US market in the aftermath of China's PNTR status acquisition.

## CONSTRUCTION OF THE VARIABLES

Using the industry specific NTR gap measure pioneered by Handley and Limao (2017), Pierce and Schott (2016) and Facchini, et al. (2019), we compute the measure of change (reduction) in trade policy uncertainty that Chinese exporters (to the US) experience from the pre-WTO regime. We measure the NTR gap for industry  $i$  as the difference between the MFN tariff rates applied by the US to the WTO members and the exorbitant tariff rates that would have been invoked if MFN status had not been renewed by the US Congress (under “column 2 tariffs of the ‘Smoot-Hawley Trade Act’”). We use the NTR gap for the year 2000 in our empirical analysis. It should be noted that in 2000, the average NTR gap (as defined below) across industries was 0.27, with a standard deviation of 0.16.

$$\text{NTR Gap}_i = \text{Non NTR Rate}_i - \text{NTRRate}_i$$

To carry out our empirical analysis, we further aggregate the industry-level NTR gap variable at the prefecture-level,<sup>9</sup> weighted by the share of exports of each industry in the export basket in 2000. As a result, the NTR gap in prefecture  $j$  is defined as,

$$\text{NTR gap}_j = \sum_i \frac{\text{Export}_{ij}}{\text{Export}_j} * \text{NTR gap}_i,$$

where  $\text{Export}_{ij}$  are the exports of industry  $i$  from prefecture  $j$  and  $\text{Export}_j$  are total exports of the prefecture  $j$ . According to the data in 2000, the average NTR gap across prefectures is 0.32, with a standard deviation of 0.09.

To compute yearly exports from China to the United States, we utilize China Custom Data (CCD) spanning from 2000 to 2007.<sup>10</sup> CCD comprises annual transaction-level data based on the Harmonized System (HS), collated by the General Administration of Customs of China. It encompasses details of each import/export transaction, including pertinent variables for our study, such as import/export classification, destination/source country, transaction value (in USD), and transaction types.

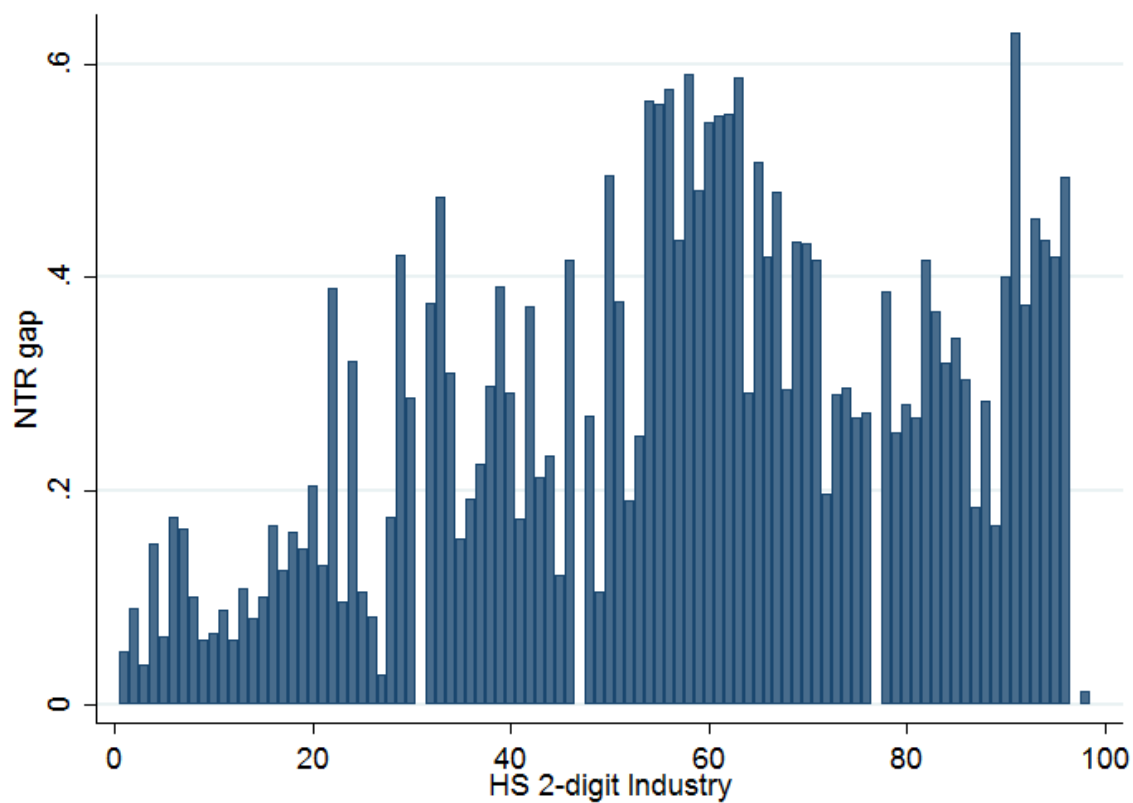
We construct the unemployment duration variable as follows. During interview, respondents recorded whether they worked for more than one hour with pay in last week. Those who were paid for at least one hour of work in the last week were then asked the exact time they started their job and how long it took them to secure the job.

We use the time these jobseekers spent in securing their jobs between January 2002 and December 2007 to measure unemployment duration. Respondents who were still unemployed at the time of the interview were asked exactly when they left their last job and whether they were actively looking for a job. In this case, we use the length of time since they had left their last job to measure unemployment duration. According to the data, 265 individuals were still unemployed at the time of the survey, accounting for 6.4% of the total samples. The definitions of all variables are provided in Table 1.<sup>11</sup>

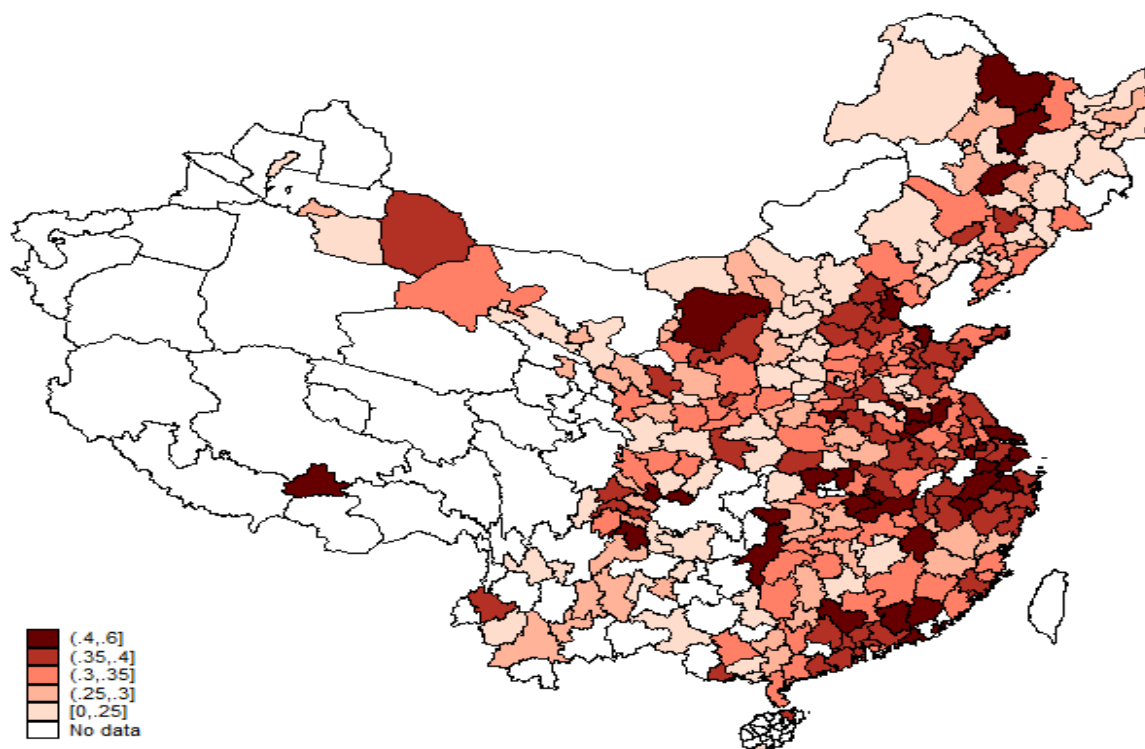
<sup>9</sup> ‘Prefectures’ are administrative subdivisions of China’s provincial-level divisions.

<sup>10</sup> We are grateful to Zhangfeng Jin (a postdoctoral student at Zhejiang University) for his help on giving us temporary access to this dataset.

<sup>11</sup> Given that we are interested in how the exogenous variations in the structural attributes of the urban jobseekers across the in-migrants and urban natives, as well as the variations in local labour demand conditions explain variation in the time it takes to end the spell of remaining unemployed owing to the trade policy change, we have chosen to define only the relevant variables in Table 1.



**Figure 1.** HS 2-Digit Industry-Level NTR Gap (2000)



**Figure 2.** Geographical Distribution of Prefecture-Level NTR Gap (2000)



**Table 1.** Description of Variables

<b>Variables</b>	<b>Definition</b>
<b>Unemployment Duration</b>	<i>The length of time elapsed between the dates of inflows and outflows from unemployment (or censoring date defined as December 2007).</i>
<b>Observation Period</b>	<i>January 2002-December 2007</i>
<b>Age</b>	<i>Years of age</i>
<b>Age<sup>2</sup></b>	<i>Age squared</i>
<b>Male</b>	<i>=1 if Male</i>
<b>College</b>	<i>=1 if holds a college degree</i>
<b>Health</b>	<i>=1 if in good health</i>
<b>Training</b>	<i>=1 if receives no job relevant training</i>
<b>Urban</b>	<i>=1 if an urban native</i>
<b>Employment Status</b>	<i>=1 if exited the unemployment spell (censoring variable)</i>
<b>Social Network</b>	<i>Size of social network of the household in log</i>
<b>Income</b>	<i>Household income in log</i>
<b>Expenditures</b>	<i>Household expenditures in log</i>
<b>Unemployment</b>	<i>Unemployment rate at the city-level</i>
<b>NTR Gap</b>	<i>Degree of reductions in tariffs on Chinese exports to the US owing to China's attainment of PNTR after its WTO entry.</i>
<b>NTR Gap×Income</b>	<i>Interaction term between NTR gap and income</i>
<b>NTR Gap×Urban</b>	<i>Interaction term between NTR gap and urban dummy</i>

### TRENDS IN UNEMPLOYMENT DURATIONS: INSIGHTS FROM DESCRIPTIVE STATISTICS

Table 2 provides the descriptive statistics for certain key variables used in the unemployment duration analysis. The mean unemployment duration of complete observations is 40 days, and 75% of those with complete observations have unemployment durations of no more than 30 days. In contrast, the mean unemployment duration of the censored observation is 807 days, and 75% of those with censored observations have unemployment durations of no fewer than 270 days. Table 3 and Figure 3 provide further evidence on the distribution of unemployment duration. In particular, Figure 3 demonstrates the probability of not finding a job over the entire duration of our study. We use the non-parametric method provided by Kaplan and Meier (1958) to estimate these probabilities. Figure 3 makes clear that more than 80% of the jobseekers find employment within 100 days.

A comparison between migrants and urban natives for the main variables used in the unemployment duration analysis is presented in Table 4. Columns 2 and 3 in Table 4 report the mean value separately for the samples of the in-migrants and urban natives. The last column tests the difference in mean value between in-migrants and urban natives for each variable. The main insights are as follows:

1. The mean unemployment durations are about 90 days, and they are much longer for urban natives (160 days) compared to the internal in-migrant jobseekers (44 days).
2. Demographically, these internal in-migrants comprise more male and younger jobseekers compared to the urban natives. However, the internal in-migrants are more likely to be in better health relative to the urban natives.

3. Internal in-migrants are characterised by (a) smaller size of social network at their place of work, (b) lower household incomes and (c) lower household expenditures compared to the urban natives.
4. The internal in-migrant jobseekers are more likely to stay in regions with lower unemployment rates compared to the urban natives.
5. Both in-migrants and urban natives in the prefectures facing average decline in trade policy uncertainty more quickly finish their cycles of unemployment.

**Table 2.** Descriptive Statistics

Unemployment Duration		
Variables	Type	Description
All	Quantitative	First Quartile: 7 Median: 20 Third Quartile: 40 Mean: 90 SD: 263
Complete	Quantitative	First Quartile: 7 Median: 15 Third Quartile: 30 Mean: 40 SD: 86
Censored	Quantitative	First Quartile: 270 Median: 605 Third Quartile: 1275 Mean: 807 SD: 641

Individual Characteristics		
Variables	Type	Description
Age	Quantitative	First Quartile: 24.00 Median: 29.00 Third Quartile: 36.00 Mean: 30.21 SD: 7.54
Male	Qualitative	2 Levels, 1: Female (48.62%) 2: Male (51.38%)
College	Qualitative	2 Levels, 1: High School and Below (62.36%) 2: College and Above (37.64%)
Health	Qualitative	2 Levels, 1: Poor (16.34%) 2: Average (83.66%)

Table 2. Continued

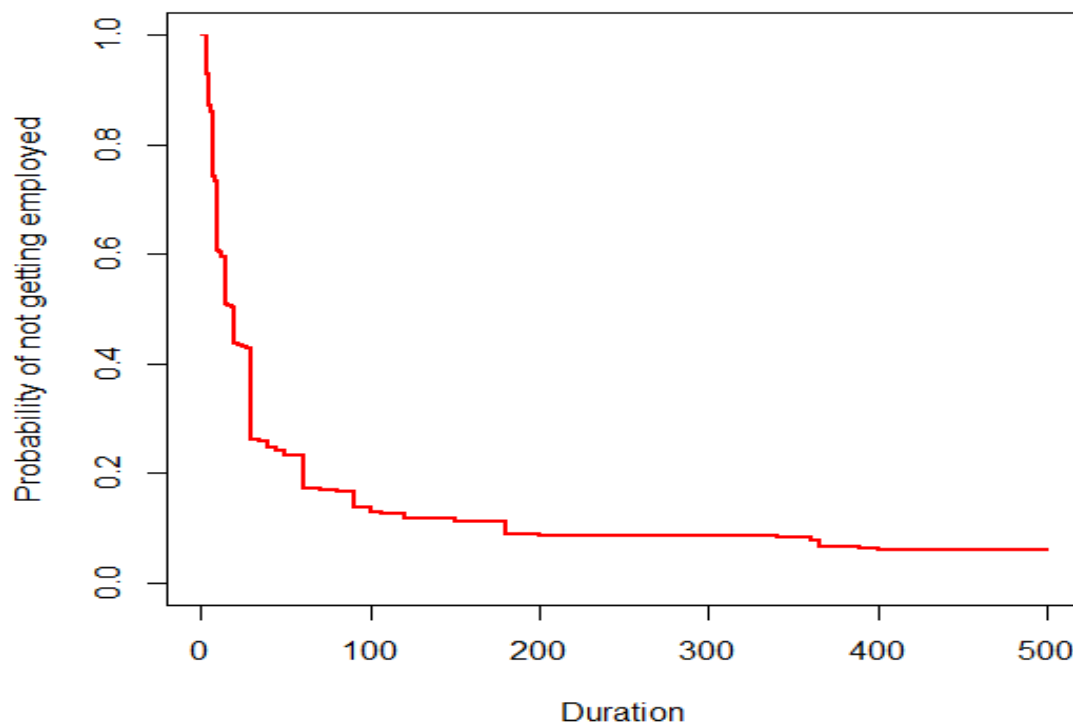
Individual Characteristics		
Variables	Type	Description
Training	Qualitative	2 Levels 1: No Training (70.43%) 2: Training (29.57%)
Urban	Qualitative	2 Levels 1: Urban Migrant (60.55%) 2: Urban Native (39.45%)
Employment Status	Qualitative	2 Levels 1: Unemployed (6.48%) 2: Employed (93.52%)
Household Characteristics		
Variables	Type	Description
Social Network	Quantitative	First Quartile: 2.40 Median: 3.05 Third Quartile: 3.50 Mean: 3.01 SD: 0.95
Income	Quantitative	First Quartile: 9.90 Median: 10.42 Third Quartile: 10.87 Mean: 10.42 SD: 0.70
Expenditure	Quantitative	First Quartile: 9.39 Median: 9.88 Third Quartile: 10.34 Mean: 9.87 SD: 0.74
Local Labour Market Conditions		
Variables	Type	Description
Unemployment	Quantitative	First Quartile: 0.04 Median: 0.08 Third Quartile: 0.11 Mean: 0.08 SD: 0.04
NTR Gap	Quantitative	First Quartile: 0.35 Median: 0.38 Third Quartile: 0.39 Mean: 0.37 SD: 0.04

**Table 3.** Frequency Distribution of Duration

Time Interval	Frequency of Complete Observations	Frequency of Censored Observations
[0,20]	2309 (60.38%)	1 (0.38%)
[20,40]	771 (20.16%)	13 (4.91%)
[40,60]	298 (7.79%)	7 (2.64%)
[60,80]	24 (0.62%)	0 (0%)
[80,100]	156 (4.08%)	15 (5.66%)
>100	266 (6.96%)	229 (86.42%)

**Table 4.** Summary Statistics for Migrants and Urban Natives

Variables	Mean (Migrants)	Mean (Urban Natives)	Diff.
Unemployment Duration	44.265	160.446	-116.181***
Age	28.949	32.142	-3.193***
Male	0.552	0.455	0.097***
Social Network (log)	2.920	3.148	-0.228***
Income (log)	10.134	10.852	-0.718***
Expenditures (log)	9.574	10.314	-0.739***
Unemployment	0.072	0.069	0.004***
NTR Gap	0.370	0.369	0.001
Observation	2,476	1,613	-

**Figure 3.** Kaplan-Meier Estimates of Unemployment Duration

## TRANSMISSION MECHANISM

The transmission mechanism between the NTR gap and unemployment duration is the following: an increase in *NTR gap* (owing to reduced trade policy uncertainty w.r.t. the border taxes to be paid by the Chinese exporters to the US) at prefecture  $j$  would induce not only an increase in the number of vacancies and stimulate demand for local labour in prefecture  $j$  but also an increase in internal in-migration to prefecture  $j$ . Hence, this should reduce unemployment duration across local urban jobseekers and the local in-migrants in the prefectures that are facing average decline in trade policy uncertainty. However, the differences across these two groups in terms of reservation income/wages and other structural factors, including attainment of *Hukou* registrations, skill-differentiation, training and age differentiations, along with heterogeneous job accessibility, should have contributed to the differential impact of the reduced trade uncertainty on the unemployment duration across these two groups of jobseekers.

Given that we have aggregated the industry-level NTR gap variable at the prefecture level (weighted by the share of exports of each industry in the export basket in 2000), the prefectures that are facing average decline in trade policy uncertainty see a shorter unemployment cycle for jobseekers.

In the following sections, we build Cox PH and AFT regression models on unemployment durations across local in-migrants and urban natives, conditional on observable characteristics.

## ECONOMETRIC FRAMEWORK: COX PH AND AFT MODELS

Duration models have been used to analyse the determinants of survival in remaining unemployed. This class of models permits the researchers to consider data specificity: the dependent variable under consideration is the time spent from the start of unemployment (unemployment spell), and observations are censored. In this case, data are right-censored, i.e., information related to the start of an unemployment span is known and, in the last moment observed, several periods have finished while others continue.

We have started with the most popular approach, viz., the Cox PH model, to deal with this type of data. Cox's PH model is formulated in terms of the hazard rate, which is the instantaneous rate of occurrence of an event at some time  $t$ , given that the event has not happened till that time. In our case, the event is indeed the end of an unemployment spell for an individual. As the name suggests, this model assumes that the hazard functions corresponding to different individuals are proportional and as such the covariates have a multiplicative effect on the hazard. Interestingly, this model does not require full specification of the hazard function. However, in addition to the Cox PH model, we have also deployed the AFT class of models to fully specify the underlying distribution of time to an event. This is precisely because the AFT models assume that the effect of a covariate is to accelerate or decelerate the time to an event, while the traditional PH models do not.

In our paper, unlike the relevant recent literature such as Caliendo, et al. (2013; 2016), we use all the relevant parametric distributions within our AFT specifications, along with the Cox PH, to represent spells of remaining unemployed. Among these various specifications, we considered the log-logistic specification which eventually turns out to be the best fit for our sample of unemployment durations data.

We start our empirical analysis by first employing the Cox PH model of unemployment duration. Suppose  $T$  represents the unemployment duration for an individual (in days) after entering the job market, where  $\mathbf{x} = (x_1, x_2, x_3, \dots)$  is a vector of variables affecting the unemployment duration for an individual. Then the Cox PH model (Cox, 1972) assumes that the variables have a multiplicative effect on the hazard and that this effect is constant over time, i.e.

$$h(t|\mathbf{x}) = h_0(t)e^{\beta_1 x_1 + \beta_2 x_2 + \dots + \beta_p x_p}, \quad (4.1)$$

where  $h_0(t)$  can be thought of as a baseline hazard function. Thus, the model parameters in this model provides the effect of the variables on the overall hazard. Since  $e^{x\beta}$  is always positive,  $h(t|\mathbf{x})$  is automatically non-negative for all  $\mathbf{x}$  and  $\beta$ . The PH model does not require explicit specification of the baseline hazard function.

As an alternative estimation strategy, we also consider another class of models, namely the AFT models (Kalbfleisch and Prentice, 1981), in order to judge the relative efficacy of different survival models for which the logarithm of durations, i.e.  $Y = \log T$ , given  $\mathbf{x}$ , has a distribution with a location parameter  $\mu(\mathbf{x})$  and a constant scale parameter  $\sigma$ . These models can be written in the following form.

$$Y = \mu(\mathbf{x}) + \sigma e, \quad (4.2)$$

where  $\mu(\mathbf{x}) = \mathbf{x}\beta$ ,  $\sigma > 0$  and  $e$  follows a distribution that does not depend on  $\mathbf{x}$ . Although the family of AFT models is distinct from the PH family, the exponential and Weibull models belong to both classes. In the AFT class of models,  $\mathbf{x}$  affects  $Y$  linearly and thereby affects  $T$  multiplicatively. We consider all the popular special cases of AFT models, the details of which are provided in Appendix (in Table A1).

Towards this end, we highlight the main difference between the PH model and AFT model. As the above model specifications suggest that a variable  $X$  has a multiplicative effect on the hazard function of the unemployment duration in the PH model, whereas it has a multiplicative effect on the unemployment duration in the AFT model. The interpretation of the PH model is typically done in terms of hazard ratios, which is defined as the ratio of the estimated hazard functions for two different values of a variable. This is possible because the ratio  $h(t|\mathbf{x}_1)/h(t|\mathbf{x}_2)$  of hazard functions for two individuals with vectors of variables  $\mathbf{x}_1$  and  $\mathbf{x}_2$  does not vary with the duration,  $t$ . A hazard ratio taking a value greater than one implies the event is more likely to occur, and a ratio less than one means an event is less likely to occur. A hazard ratio that equals to one means the variable has no effect on the hazard of the event. Thus, intuitively, it is easy to see that the time to an event  $T$  increases if the corresponding regression coefficient in the PH model is negative and vice versa.

## RESULTS

In this section, we provide the results for both Cox PH and AFT models. We use all four special cases for AFT models, i.e., Exponential, Weibull, Log-normal and Log-logistic, as summarised in Appendix Table A1. The log-logistic AFT model is found to be the best-fit model among the fitted models (the last column in Table 5).<sup>12</sup> It is worthwhile to mention that the sign of all the coefficients for the main variables of interest have reversed for AFT models as compared to the PH model in Table 5. As discussed in Section 4, this is primarily because of the difference in specifications of these two classes of models. Furthermore, this also shows the robustness of our results as the effect of the variables remain consistent across all these models.<sup>13</sup>

We use reductions in trade policy uncertainty on exports from China to United States, proxied by the reduction in tariffs imposed by the US on Chinese exports. This is measured by the variable NTR Gap (defined in Table 1) as a proxy for a positive demand shock on the local labour market. In line with our baseline estimation results based on the log-logistic AFT model, we find that a larger demand shock in the exportable sector of the local labour market (measured by NTR Gap) is associated with shorter unemployment duration and is statistically significant. Note that the corresponding coefficient

<sup>12</sup> This is based on the information criteria (AIC, SBC).

<sup>13</sup> See next page footnote.

in Cox PH model is positive, which is consistent with our discussion in Section 4. This suggests that the unemployment duration is more likely to get longer as the NTR Gap reduces.

Looking at the coefficient corresponding to the “Income” variable in Table 5, we can deduce that household income is negatively associated with individuals’ unemployment duration, while a relatively higher household expenditure is associated with longer unemployment duration. The corresponding PH model coefficient is also consistent with this observation., which can also be explained with the help of the *job accessibility hypothesis*. Individual labourers belonging to a more (less) wealthy households are likely to have more (lower) job options and are more (less) likely to exit the unemployment pool. It is also possible that individual jobseekers from more (less) wealthy households can subject themselves to higher (lower) reservation wages, and in turn can remain unemployed until they find job that meets their own standards and preferences and offers acceptable remuneration for their services. Our result thus confirms the relevance of job accessibility channels over the reservation wage hypothesis.

Using the local unemployment rate as a proxy for local labour market conditions, we find that a higher unemployment rate is associated with longer unemployment duration. This could be because a higher unemployment rate is associated with lower labour demand relative to labour supply, implying that it is more difficult for individuals to find a job locally.

Results in Table 5 suggest that with greater ease of internal migration, a migrant jobseeker has a much shorter unemployment duration period relative to urban natives. The corresponding PH model coefficient also suggests that a migrant jobseeker is more likely to find a job as compared to an urban native. There could be at least two different theoretical explanations. First, internal in-migrants are more underprivileged in the urban labour market space due to the *Hukou* system in China. In a sense, urban natives will have greater priority in access to the local labour market than working in different locations, assuming a comparable level of skill and wages in both locations. Indeed, as emphasized by Chacko and Chelminski (1996), urban local individuals who benefit from housing perks in the location of their current employment tend to be hesitant to pursue new opportunities, even when offering higher compensation. As a result, in-migrant jobseekers are likely to hold down their reservation wages relative to jobseekers among urban natives. In this case, lower reservation wages lead to shorter an unemployment spell for migrant jobseekers. An alternative reason could be the lower mobility costs of the migrant jobseekers across sectors, which could make them more sensitive to job opportunities relative to urban natives. In the latter case, higher job accessibility accounts for shorter unemployment duration.

While both these channels can give rise to a negative association between migration and unemployment duration, the policy implications can prove quite different. If migration reduces unemployment duration through lower reservation wage, this does not necessarily imply that migration is welfare-improving, particularly when migration is associated with lower reemployed

<sup>13</sup> For an illustration, let us consider the Weibull regression model which is a special case of both PH and AFT class of models. Given a predictor  $X$ , the AFT model is given by

$$\log(T) = \mu(x) + \sigma e,$$

where  $\mu(x) = \log(\eta(x)) = \beta_0 + \beta_1 x$  and  $\sigma = 1/\delta$ , with  $\eta(x)$  and  $\delta$  being the scale and shape parameters of the Weibull distribution, respectively.

It is known that the hazard function in this case is given by

$$\begin{aligned} h(t|x) &= \left(\frac{\delta}{\eta(x)}\right) \left(\frac{t}{\eta(x)}\right)^{\delta-1} \\ &= h_0(t) \exp[-\delta(\beta_0 + \beta_1 x)], \end{aligned}$$

with  $h_0(t) = \delta t^{\delta-1}$  being the baseline hazard rate. Since  $\delta > 0$ , there is a sign reversal in this case.

wages. In contrast, if migration reduces unemployment duration through higher job accessibility, we can argue more convincingly that migration is resulting in the reduction of spatial mismatch and is therefore welfare-improving.

**Table 5.** Model Summary for Cox and AFT Regression Models (without any Interaction Effects)

Coefficients	Cox	Exponential	Weibull	Log-Normal	Log-Logistic
<b>Constant</b>	-	2.239*** (0.394)	2.100*** (0.608)	1.608*** (0.542)	1.449*** (0.504)
<b>Age</b>	-0.095*** (0.019)	0.140*** (0.019)	0.149*** (0.029)	0.110*** (0.026)	0.096*** (0.024)
<b>Age<sup>2</sup></b>	0.001*** (0.000)	-0.002*** (0.000)	-0.002*** (0.001)	-0.002*** (0.000)	-0.001*** (0.000)
<b>Male</b>	0.138*** (0.033)	-0.537*** (0.034)	-0.474*** (0.051)	-0.185*** (0.044)	-0.077* (0.041)
<b>College</b>	0.142*** (0.050)	-0.238*** (0.048)	-0.270*** (0.075)	-0.164** (0.067)	-0.132** (0.064)
<b>Health</b>	0.164*** (0.045)	-0.242*** (0.045)	-0.278*** (0.069)	-0.225*** (0.060)	-0.214*** (0.058)
<b>Training</b>	0.173*** (0.036)	-0.366*** (0.036)	-0.329*** (0.056)	-0.217*** (0.049)	-0.184*** (0.046)
<b>Urban</b>	-0.639*** (0.053)	1.257*** (0.051)	1.319*** (0.080)	0.881*** (0.070)	0.763*** (0.068)
<b>Social Network</b>	0.050*** (0.017)	-0.006 (0.017)	-0.033 (0.027)	-0.076** (0.024)	-0.090*** (0.023)
<b>Unemployment</b>	-0.870* (0.464)	0.636 (0.474)	0.575 (0.718)	0.800 (0.609)	0.991* (0.566)
<b>NTR Gap</b>	2.030*** (0.480)	-1.787*** (0.467)	-2.155*** (0.732)	-2.389 (0.654)	-2.660*** (0.613)
<b>Income</b>	0.138*** (0.039)	-0.346*** (0.040)	-0.323*** (0.061)	-0.132** (0.054)	-0.087* (0.050)
<b>Expenditures</b>	-0.252*** (0.037)	0.441*** (0.038)	0.392*** (0.058)	0.228*** (0.051)	0.212*** (0.047)
<b>Year [2003]</b>	-0.097 (0.073)	-0.152** (0.074)	-0.103 (0.113)	0.024 (0.098)	0.099 (0.093)
<b>Year [2004]</b>	-0.107 (0.073)	-0.049 (0.074)	-0.001 (0.113)	0.046 (0.098)	0.056 (0.093)
<b>Year [2005]</b>	0.007 (0.069)	-0.376*** (0.070)	-0.263** (0.107)	-0.098 (0.092)	-0.055 (0.087)
<b>Year [2006]</b>	-0.044 (0.067)	-0.462*** (0.068)	-0.289*** (0.104)	-0.063 (0.090)	-0.020 (0.084)
<b>Year [2007]</b>	-0.039 (0.065)	-0.635*** (0.067)	-0.400*** (0.101)	-0.081 (0.087)	0.011 (0.081)
<b>AIC</b>	56738.76	39595.35	37728.76	36250.35	36000.32

**Note:** \*\*\*:  $p$  – value  $< 0.01$ , \*\*:  $p$  – value  $< 0.05$  and \*:  $p$  – value  $< 0.1$ . Standard errors are in parentheses. The results comprise year dummies measuring when the spell of unemployment duration was ended within 2002-2007 timespan, viz., when the respective respondent has started the present job.

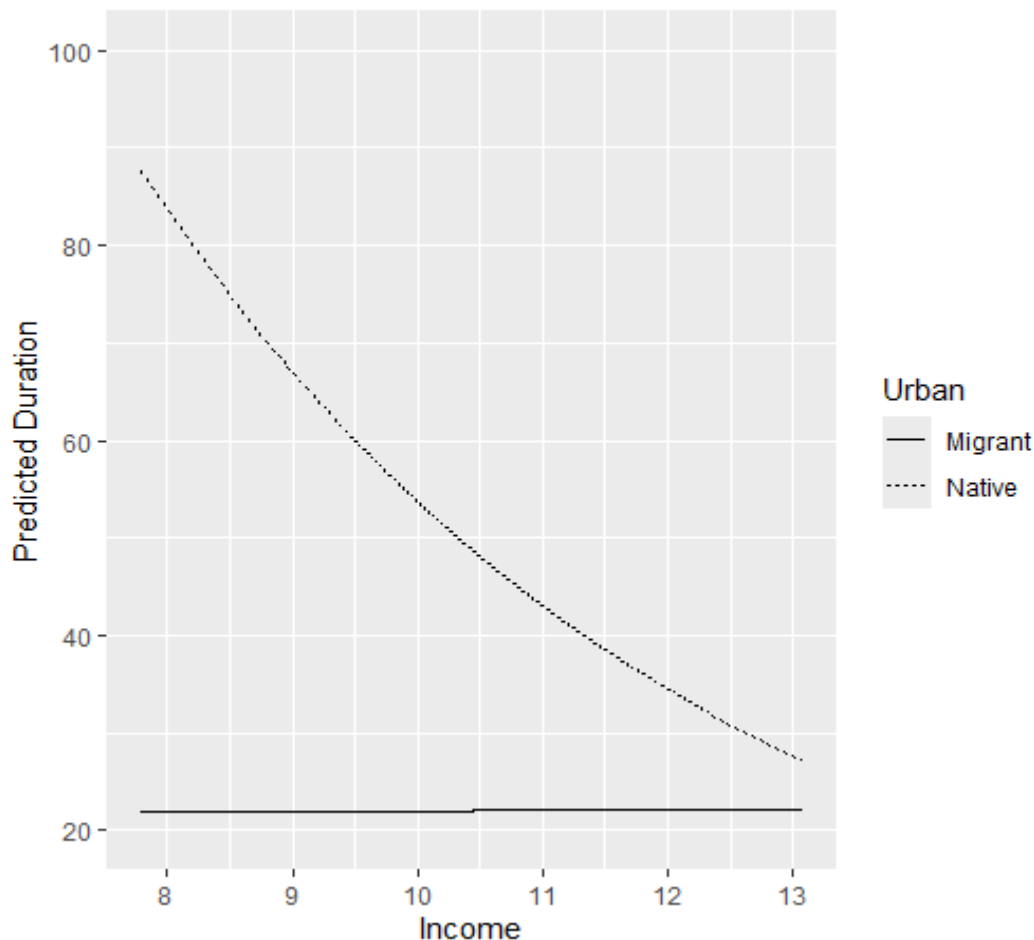


To test the two above-mentioned channels, viz., heterogeneity in job accessibility and reservation wages across the local in-migrants and urban natives, we have re-estimated all the duration models by adding the two interaction dummies:

1. between migration (the dummy variable “Urban”) and household income (the variable “Income”);
2. between migration dummy and extent of reductions in tariffs imposed by the US on Chinese exports to the US, owing to China’s attainment of the PNTR status (this interaction term is captured by the term “NTR gap×Urban”).

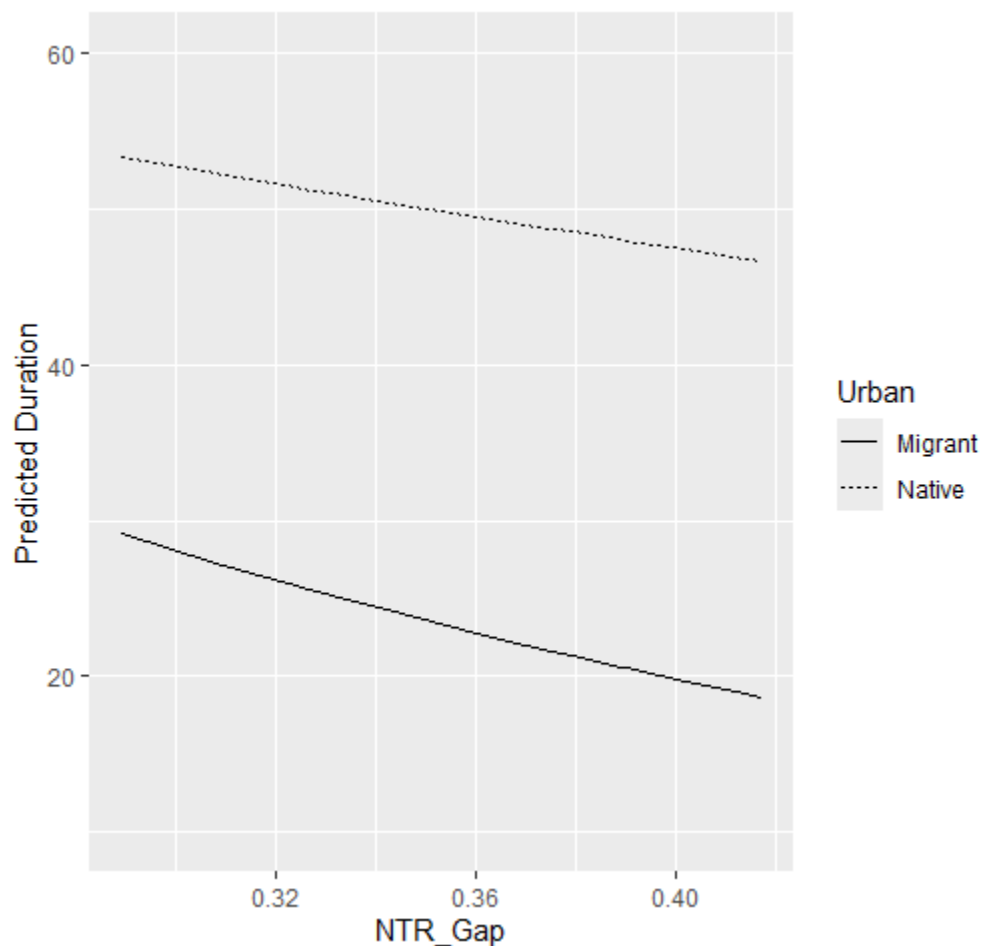
The results of the models with these two interaction dummies are presented in Table 6. As our best fitted model is the log-logistic AFT model for unemployment duration, we report the results for the coefficients of the interaction dummies in the last column of Table 6.

The coefficient of the interaction term (Income×Urban) indicates whether the migrant-urban gap in unemployment duration responds to household income. The estimated result reveals that the migrant-urban gap in unemployment duration is greater for individuals from households with lower incomes. Given the fact that lower-income households are associated with lower job accessibility in the local labour market, this finding indirectly supports the fact that migrants’ lower unemployment duration can partially be explained by higher job accessibility outside their local labour markets.



**Figure 4.** Interaction Effect between Income and Urban Dummy Variable

The coefficient of the interaction term (Income $\times$ Urban) indicates how the migrant-native gap in unemployment duration responds to household income. This can be easily understood with the interaction plot in Fig. 4, which shows how the predicted unemployment duration changes as the household income increases both for Urban natives and migrants. Note that this plot is created after fixing other continuous predictors in the model at their respective average values and the qualitative predictors at their reference levels. It is easy to understand that the migrant-native gap in unemployment duration is greater for individuals from households with lower incomes. Given the fact that lower-income households are associated with lower job accessibility in the local labour market, this finding indirectly supports the fact that migrants' lower unemployment duration can partially be explained by higher job accessibility outside their local labour markets.



**Figure 5.** Interaction Effect between NTR Gap and Urban Dummy Variable

Next, we focus on the interaction between 'NTR gap' and 'Urban' dummy variable. Fig. 5 shows how predicted unemployment duration changes with NTR gap both for Urban natives and migrants. We find that the migrant-native gap in unemployment duration is larger as NTR Gap increases. Note that NTR Gap is higher in the regions that are exposed to the positive demand shocks in the export markets owing to China's attainment of the PNTR status after its accession to the WTO. Given that such export shocks should result in more labour demand, our finding suggests that migration increases job accessibility which in turn reduces the unemployment duration for both internal in-migrants and urban natives.

**Table 6.** Model Summary for Cox and AFT Regression Models (with Interactions between (i) NTR Gap and Urban and (ii) NTR Gap and Income)

Coefficients	Cox	Exponential	Weibull	Log-Normal	Log-Logistic
Constant	-	0.927** (0.460)	0.862 (0.703)	0.935 (0.632)	0.991* (0.576)
Age	-0.091*** (0.019)	0.136*** (0.019)	0.141*** (0.029)	0.103*** (0.026)	0.090*** (0.024)
Age <sup>2</sup>	0.001*** (0.000)	-0.002*** (0.000)	-0.002*** (0.000)	-0.001*** (0.000)	-0.001*** (0.000)
Male	0.135*** (0.033)	-0.534*** (0.034)	-0.469*** (0.051)	-0.179*** (0.044)	-0.074* (0.041)
College	0.125* (0.051)	-0.213*** (0.048)	-0.240*** (0.075)	-0.142** (0.068)	-0.117* (0.064)
Health	0.155*** (0.045)	-0.221*** (0.045)	-0.257*** (0.070)	-0.214*** (0.060)	-0.205*** (0.058)
Training	0.168*** (0.037)	-0.363*** (0.037)	-0.322*** (0.056)	-0.211*** (0.050)	-0.180*** (0.046)
Urban	-2.333*** (0.646)	4.448*** (0.641)	4.580*** (0.990)	2.729*** (0.863)	2.239*** (0.827)
Social Network	0.046*** (0.017)	0.002 (0.017)	-0.024 (0.027)	-0.068*** (0.024)	-0.084*** (0.023)
Unemployment	-0.760 (0.465)	0.502 (0.472)	0.385 (0.714)	0.637 (0.610)	0.827 (0.568)
NTR Gap	2.722**** (0.595)	-3.155*** (0.575)	-3.487*** (0.901)	-3.391*** (0.829)	-3.457*** (0.756)
Income	0.038 (0.046)	-0.142*** (0.048)	-0.124* (0.072)	-0.009 (0.064)	0.003 (0.058)
Expenditures	-0.244*** (0.037)	0.417*** (0.038)	0.371** (0.058)	0.218*** (0.051)	0.205*** (0.047)
NTR	-2.013* (0.977)	3.808*** (0.947)	3.818** (1.484)	2.724** (1.317)	2.423* (1.268)
Gap×Urban	0.231*** (0.056)	-0.436*** (0.056)	-0.443*** (0.086)	-0.271*** (0.077)	-0.224*** (0.073)
Income×Urban	-0.097 (0.073)	-0.155** (0.074)	-0.106 (0.113)	0.021 (0.098)	0.097 (0.093)
Year [2003]	-0.100 (0.073)	-0.029 (0.074)	0.001 (0.113)	0.035 (0.098)	0.050 (0.093)
Year [2004]	0.006 (0.069)	-0.352*** (0.070)	-0.252** (0.106)	-0.107 (0.092)	-0.062 (0.087)
Year [2005]	-0.046 (0.067)	-0.435 (0.068)	-0.273*** (0.104)	-0.065 (0.090)	-0.021 (0.084)
Year [2006]	-0.040 (0.065)	-0.603*** (0.067)	-0.386*** (0.101)	-0.087 (0.087)	0.006 (0.081)
Year [2007]					
AIC	56738.76	39529.29	37702.7	36239.9	35993.26

Notes: \*\*\*: p-value<0.01, \*\*: p-value<0.05 and \*: p-value<0.1. Standard errors are in parentheses. The results comprise year dummies measuring when the spell of unemployment duration was ended within 2002-2007 timespan, viz., when the respective respondent has started the present job.

Furthermore, as revealed in Table 5 and Table 6, the results regarding the individual factors are also found to be quite significant in determining unemployment durations among urban locals and internal migrants. For instance, *age* is positively associated with unemployment durations, implying that older jobseekers are at greater disadvantage in finding a job. This is consistent with the fact that in Chinese labour-intensive industries, younger jobseekers who are not necessarily considered less competent, remain dominant. The *negative coefficient of age-squared* reveals that unemployment duration is not a monotonically increasing function of age. To be precise, after reaching a threshold (about 34 years old), unemployment duration is found to decrease with age. This inverted U-shaped association between unemployment duration and age reflects the fact that as individuals get older, they are likely to accumulate more skills in a specific field, which also increases their chance of being re-employed in our sample.<sup>14</sup>

The last columns of Table 5 and Table 6 also reveal that *male individuals are more likely to be re-employed than female individuals*, which describes the role of gender discrimination in the Chinese labour market. This characteristic feature is more common in the developing economies than in the advanced economies. For example, female individuals are not provided the same job opportunities as their male counterparts, which makes it more difficult for females to find jobs. An alternative explanation is that male individuals make more effort to find jobs, given that they shoulder greater responsibility as the primary earners in traditional Chinese families.

Individuals who have acquired a college degree and beyond are more likely to exit the unemployment pool than other individuals with lower education qualifications. The higher exit rate of more educated individuals can be explained by their higher capacity to search for jobs through better access to information and wider ranges of alternatives for future employment. We further find that a better health condition, as an alternative measurement of human capital, is associated with shorter spells of unemployment duration. Receiving job-relevant training has a negative impact on unemployment duration. One explanation is that job training can be considered as an investment in specific human capital, which is complementary to general human capital such as education and health in search of a job.

Individuals exposed to a larger-sized household social network are associated with shorter unemployment duration. A possible explanation is that a larger size of social network will facilitate more job opportunities through better access to job information, which is consistent with the fact that the social network has proven an important factor in determining labour market outcomes.

Overall, this study comes with significant evidence that *job accessibility and income differences* are important channels to explain the differences in unemployment duration, owing to the local labour demand shocks induced by the reduction in uncertainty regarding the lower MFN tariff rates to be faced by Chinese exports in the US market. In other words, the internal in-migrants find it relatively easier to finish their cycles of remaining unemployed, relative to the case of the urban natives who are inflexible about searching for jobs beyond their locality and are also unwilling to accept wages below their reservation wages.

## CONCLUSIONS AND FUTURE RESEARCH AGENDA

Reductions in policy uncertainty relating to the exporting of goods from China to the US, owing to China's accession to the World Trade Organisation (WTO), affect the demand for labour within the domestic labour market and in turn impact the internal relocation of prospective workers (i.e.,

<sup>14</sup> The readers may draw an inference from the above results that is consistent with general employment dynamics observed in the real world. Young entrants into the job markets are preferred more as starters of new jobs compared to the relatively senior jobseekers, given that the young entrants are likely to be more productive and likely to demand lower wages compared to senior jobseekers. However, the presence of experienced senior colleagues within any organization is also essential to improving the productivity of the new entrants, thus retaining or re-employing experienced applicants is likely.

jobseekers) in terms of their relative durations of remaining unemployed. We exploit the survey on Rural Urban Migration in China (RUMiC), conducted in early 2008, to construct an inflow sample of the unemployed, made up of those who started a period of joblessness between January 2002 and December 2007. This allowed us to empirically determine the reasons behind heterogeneity in unemployment durations among local jobseekers in urban China in the post-WTO regime. Using the classes of PH and AFT models, we show how the positive boost in exports, following the reduction in trade policy uncertainty regarding Chinese exports to be sold to the US market, and owing to the attainment of PNTR status, would affect the reallocation of labour, income (real wages) and employment status for the rural migrants and the urban natives within China's internal labour market. Overall, we find robust evidence that the reductions in trade policy uncertainty significantly shorten urban jobseekers' unemployment duration. However, the migrant jobseekers continue to face relatively shorter unemployment durations compared to local natives, particularly in those regions that are exposed to the positive demand shocks in the export markets owing to China's PNTR status. The individual-specific factors such as age, education, experience, health status, gender, household income and the relative size of the social network have also been declared prominent in exaggerating the heterogeneity in unemployment durations across urban jobseekers.

In the light of these findings, it would be safe to argue that the deregulation of the *Hukou* system did have a positive impact in enhancing the migration within China. Therefore, such structural reform policy should be continued to homogenise the internal labour market. There are several ways to extend our study in the future.

First, given our limitations of information about unemployment spells beyond December 2007, we could not explore the joint implications of reduced trade policy uncertainty induced demand shock and that of the introduction of 2008 *Labour Contract Law* upon the unemployment durations across the non-*Hukou* in-migrants and urban natives. Also, with such an extensive sample beyond 2008, one could compare the unemployment durations against the counterfactual scenario where China would have completely relaxed the *Hukou* system, enabling all the prospective in-migrants to be relocated to, and work in, urban sectors, being at par with the urban natives, in making their living. Second, one can examine whether there will be an oversupply of migrant jobseekers relative to the urban natives, even in the labour-intensive sectors that are closely associated to the PNTR status. Moreover, while it would be interesting to consider if the *Hukou* system is completely deregulated, would slums emerge in Chinese cities like what happened in India? This can be another useful direction of future research.

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## APPENDIX

For a relative comparison of all the parametric models in our unemployment duration model, Table A1 is produced below.

**Table A1.** Parametric Models for Unemployment Duration

Name	$f(t)$	$S(t)$	$G(z)$	$\sigma$
Exponential	$\frac{1}{\eta} e^{-\left(\frac{t}{\eta}\right)}$	$e^{-\left(\frac{t}{\eta}\right)}$	$\exp(-e^z)$	1
Weibull	$\left(\frac{\beta}{\eta}\right) \left(\frac{t}{\eta}\right)^{\beta-1} e^{-\left(\frac{t}{\eta}\right)^\beta}$	$e^{-\left(\frac{t}{\eta}\right)^\beta}$	$\exp(-e^z)$	$\frac{1}{\beta}$
Log-normal	$\frac{1}{\sqrt{2\pi\sigma^2}t} e^{-\frac{1}{2}\left(\frac{\log t - \mu}{\sigma}\right)^2}$	$1 - \Phi\left(\frac{\log t - \mu}{\sigma}\right)$	$1 - \Phi(z)$	$\sigma$
Log-logistic	$\frac{\beta}{\alpha} \frac{(t/\alpha)^{\beta-1}}{\left[1 + (t/\alpha)^\beta\right]^2}$	$\frac{1}{1 + (t/\alpha)^\beta}$	$\frac{1}{1 + e^{-z}}$	$\frac{1}{\beta}$