THESIS ON EXCHANGE RATE AND TRADE REFORMS

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DECLARATION

I hereby declare that this thesis is my original work and it has been written by me in its entirety. I have duly acknowledged all the sources of information which have been used in the thesis.

This thesis has also not been submitted for any degree in any university previously.

Zhou Yingke
August 16 2014

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Summary

This thesis consists of three chapters within the broad field of international trade. All three essays are self-contained and can be read independently of the others. They include: (i) Do Exports Respond to Exchange Rate Changes? Inference from China's Exchange Rate Reform; (ii) Exchange Rates and Export Structure; and (iii) When Trade Discourages Political Favoritism: Evidence from China.

The first chapter revisits the exchange rate disconnect puzzle by using monthly data and exploiting the unexpected exchange rate reform in China as a natural experiment. The difference-in-differences estimation uncovers a negative and statistically significant effect of a currency appreciation on exports: a 1% currency appreciation is found to cause total exports to fall by 1.61%. We find no trade diversion by Chinese exporters after the currency appreciation, both intensive-margin and extensive margin effects of exchange rate changes on exports, and heterogeneous effects across regions, firms and industries/products.

The second chapter studies whether changes in the exchange rate affect a country's export structure, using an arguably exogenous sudden appreciation of renminbi on July 21, 2005 as the main source of identification. Employing combined regression discontinuity and difference-in-differences approach, we show that China's export structure became more similar to that of the developed countries after the currency appreciation. We also find that the

majority of the appreciation effect comes from the inter-firm resource reallocation rather than the inter-region or intra-firm resource reallocation.

The last chapter empirically investigates if trade leads to market reallocation away from politically favored but unproductive firms. This paper finds that tariff reductions after China's WTO accession induced a 2.5% percentage-point decline in the SOE output share and reduced the standard deviation of firm productivity by 1.4% between 2001 and 2005. The likelihood of SOE exit was related more to political affiliation than to performance: the SOEs affiliated with county and township governments were the worst hit, while those affiliated with higher-level governments were barely affected. Our results suggest that trade could help reduce inefficiencies arising from the political economy.

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Chapter One

Do Exports Respond to Exchange Rate Changes? Inference from China's Exchange Rate Reform

1.1 Introduction

"Japanese exports could be badly hurt by the yen's recent rapid rise, Mr. Gaishi Hiraiwa, chairman of the Keidanren, the country's federation of economic organizations, warned yesterday ..." – Financial Times, September 29 1992[1]

"In a weekend interview, Finance Minister Guido Mantega stated flatly that Brazil 'will not let the real appreciate.' A strong Brazilian real, Mr. Mantega said, hurts exports and manufacturers" – The Wall Street Journal, September 20 2012[2]

Government officials and commercial circles across the world are concerned about the severe consequences of a currency appreciation on exports and domestic production, as exemplified by the above quotes. However, academic studies show that the exchange rate movement is largely disconnected from fundamentals such as exports (this is referred to as the *exchange rate disconnect puzzle*. See Obstfeld and Rogoff, 2000).[3] For example, Dekle, Jeong and Ryoo (2010) find that the elasticity of exports with respect to the exchange rate is not statistically different from zero for every

[1] See "Japanese fear rising yen will hurt exports" by Financial Times (http://www.lexisnexis.com.libproxy1.nus.edu.sg/ap/academic/) Access date: October 9 2012

[2] See "Brazil Faces Currency Appreciation After Fed Move -Bradesco" by The Wall Street Journal (http://online.wsj.com/article/BT-CO-20120920-709858.html) Access date: October 9 2012

[3] Papers linking import prices to exchange rates include Goldberg and Knetter (1997) and Campa and Goldberg (2005, 2010), among others.

G-7 country for the period of 1982–1997.[4] The contrasting views between political/commercial circles and academia present an interesting research question: do exports respond to exchange rate changes?

Our study contributes to the aforementioned debate by revisiting the empirical evidence in two new manners. Firstly, in contrast to the yearly data that are commonly used in the literature, our empirical analysis uses monthly data, which gives us more variations with which to calculate the effect of exchange rate changes on exports. Secondly, and more importantly, instead of resorting to a micro-level analysis (i.e., using firm-destination or firm-product-destination data) as in some of the recently emerged literature,[5] we stick to the macro-level analysis but explore a natural experiment setting in China to carefully address the estimation biases due to the endogeneity associated with exchange rate changes (i.e., omitted variables bias and reverse causality).[6] Specifically, the Chinese government unexpectedly revalued its currency against the US dollar on July 21, 2005, which resulted in an immediate appreciation of 2.1% (for more description of the episode, see Section 1.3). Such an exogenous shock provides us with an opportunity to consistently estimate the effect of exchange rate changes on exports by comparing China's monthly exports to the U.S. (the treatment group) with

[4] See also Kenen and Rodrik (1986), Hooper, Johnson and Marquez (2000), and Colacelli (2009) for similar findings.

[5] See for example, Dekle, Jeong, and Ryoo, (2010); Berman, Martin, and Mayer (2012); Amiti, Itskhoki and Konings (2014); Chatterjee, Dix-Carneiro, and Vichyanond (2013).

[6] Understanding the aggregate-level response is important for both policy and academic purpose. Firstly, whether total exports respond to exchange rate movement or not is what concerns policy makers and its answer has implication for other monetary policies like interest rate, current account management, etc. Secondly, as one of the major puzzles in international macroeconomics, the small elasticity of export to exchange rate has generated a vast number of studies to understand the underlying reasons and to evaluate the potential welfare impacts of related policies (e.g., Duarte, 2003). However, there is still no consensus regarding the empirical association between exchange rate changes and total exports, due to the prevailing endogeneity issues.

those to other countries (the control group) before and after the currency revaluation, or a difference-indifferences estimation specification. Meanwhile, we also control for those potential omitted variables implied by the micro-level analysis, such as producer dispersion (Dekle, Jeong, and Ryoo, 2010; Berman, Martin, and Mayer, 2012) and import value (Amiti, Itskhoki and Konings, 2014).

We find a negative and statistically significant effect of a currency appreciation on exports. In terms of economic magnitude, a 1% currency appreciation is found to cause total exports to fall by 1.61%. Given that China exported US$1.904 trillion worth of goods in 2011, a 1% currency appreciation means a US$30.65 billion decrease in Chinese exports to the U.S., a significant number, which may justify the concerns by government officials and exporters. Our estimation results are robust to various checks on the validity of the DID estimation, including the control for country-specific month effects and country-specific linear time trend, a check on the pre-treatment differential trends between the treatment and control groups, a placebo test using homogeneous goods as the regression sample, and a difference-in-difference-in-differences (triple difference) estimation. Meanwhile, we find that the currency appreciation did not lead to trade diversion to other countries by Chinese exporters, suggesting that the fall in exports resulted in substantial exits of Chinese exporters from the exporting market. Moreover, we find the export response to exchange rate changes to be more prominent in China's coastal regions, among Chinese state-owned enterprises, within time sensitive industries, and for non-necessities.

To understand how exchange rate changes affect exports, we extend the Melitz and Ottaviano (2008) model to incorporate the role of exchange rate movements (See the Appendix 1 for details). It is found that the effect of exchange rate changes on the aggregate export value can be decomposed into two parts, the intensive and the extensive margins. Specifically, the currency appreciation increases the final prices of exports in foreign markets as well as decreases the free on board (FOB) export price due to incomplete pass-through, which causes FOB export revenues to fall (the intensive-margin effect). In the meantime, as exporters differ in their production efficiency, some less productive exporters find that their export profits become negative and hence choose to exit the foreign markets (the extensive-margin effect). By exploring our comprehensive data, we find supports for both intensive-margin and extensive-margin effects, that is, fewer firms export and for continuing exporters, each exports less, after a currency appreciation.

In addition to the aforementioned macro-level literature on the exchange rate puzzle, our study is related to recent studies using firm-level data to examine the effect of exchange rate changes on exports. For example, Dekle, Jeong, and Ryoo (2010) use panel data of Japanese exporters for the period of 1982-1997 and find the exchange-rate elasticity of exports to be statistically significant and have a value of -0.77. Drawing on French firm-level data for the period of 1995-2005, Berman, Martin, and Mayer (2012) uncover the heterogeneous reaction of exporters to real exchange rate changes: high-performance exporters increase their markup but reduce their export volume in response to a currency depreciation. Amiti, Itskhoki and Konings (2014), using Belgian firm-product level data, uncover that larger exporters also import a large amount of intermediate inputs, thereby offsetting exchange rate effects on their marginal costs and explaining the low pass-through of exchange rate changes. Chatterjee, Dix-Carneiro, and Vichyanond (2013) study the effect of exchange rate shocks on export behavior (including the adjustments of prices, quantities, product scope, and sales distribution across products) of multi-product firms. The departure of our work from these studies is that firstly we look at the aggregate export response as those in the previous literature on the exchange rate disconnect puzzle, and secondly we use a quasi-natural experiment setting to carefully control for the endogeneity problems.

Our work is also related to the literature on China's exchange rate movement. Using the same data as ours, Tang and Zhang (2012) find a significant effect of exchange rate appreciation on the exit and entry of Chinese exporters as well as on product churning. Li, Ma, Xu (2013) use detailed Chinese firm-level data to examine the effect of exchange rate changes on firms' exporting behavior, such as export volume, export price, the probability of exporting, and product scope. The main difference between our work and this literature lies in the identification strategy: while we explore the currency revaluation in July 2005 as an exogenous variation, these papers mostly rely on the panel fixed-effect estimation.

1.2 Estimation Strategy

1.2.1 Data

Our study draws on data from two sources. The first one is the China customs data from 2000 (the earliest year of the data) to 2006 (the most recent year the authors have access to). This data set covers a universe of all monthly import and export transactions by Chinese exporters and importers, specifically including product information (HS 8-digit level classification), trade value, identity of Chinese importers and exporters, and import and export destinations. The second data source is the International Financial Statistics (IFS) maintained by the International Monetary Fund (IMF), from which we obtain the monthly bilateral nominal exchange rates between China and other foreign countries as well as CPIs for the 2000-2006 period.

After combining the China customs data with the IFS data and excluding countries without monthly export value, import value and nominal exchange rate, we end up with a total of 88 countries. We then go through a few steps of data cleaning. First, we exclude 30 countries (including 9 oil-producing countries) whose currencies were pegged to the U.S. dollar in some years during our sample but unpegged in other years (see Obstfeld and Rogoff, 1995, for the same practice). Second, we exclude Hong Kong and Macao, which are largely trading centers for Chinese exports (i.e., re-export a lot of their imports from China).[7]

Table 1.1 lists the 56 countries used in our regression analysis. During our sample period, these 56 countries capture the majority of Chinese total exports, i.e., around 70%. However, one may be concerned that the revaluation of the Chinese currency coincides with a large share of exports going to countries other than those covered in the regression analysis, which would lead to an overestimation of the effect of the exchange rate change. To check such a possibility, we ploy in Figure 1.1 the share of Chinese total

[7] Results including these two economies remain qualitatively the same (available upon request).

exports covered in our regression analysis by month throughout the sample period. It is found that this number hooves around 70% and more importantly, there is no discontinuity at the time of the revaluation of the Chinese currency. These findings largely dispel the concern that our estimates may be biased due to a trade diversion to countries out of our regression sample.

Of 55 non-U.S. countries, none has its currency pegged to the U.S. dollar. Hence, we have one treatment country, the U.S., and 55 countries in the control group. Our final regression sample contains 56 × 84=4,704 country-month observations.

1.2.2 China's Exchange Rate Reform in July 2005

Timeline. After the financial crackdown in 1994, China adopted a decade-old fixed exchange rate regime, in which its currency (*RMB*) was pegged to the U.S. dollar at an exchange rate of 8.28. At 19:00 of July 21, 2005 (Beijing time), the People's Bank of China (PBOC, the central bank of China) suddenly announced a revaluation of the Chinese currency against the U.S. dollar, which was set to be traded at an exchange rate of 8.11 immediately, i.e., an appreciation of about 2.1%. Meanwhile, the PBOC announced its abandonment of the fixed exchange rate regime and that it would allow RMB to be traded flexibly with a *reference basket* of currencies with the target for RMB set by the PBOC every day. Figure 1.2 displays the trends of exchange rates of the U.S. dollar and other currencies against Renminbi during 2000-2006 (see Table 1.1 for the 55 other countries used in the analysis).[8] It is

[8] The pairwise correlation between the Renminbi-US dollar exchange rate and Renminbi exchange rate for each of the 55 countries in the control set for the post-reform period: mean value -0.03.

clear that there was a sudden drop in the exchange rate of the Chinese currency against US dollar in July 2005, and a steady and continuous decrease after that. By the end of 2006, the Renminbi had appreciated by about 5.5% against the US dollar. In the meantime, after a period of two years depreciation, the Renminbi remained quite stable against other currencies between 2004 and 2006.

Exogeneity. Despite the fact that the revaluation of the Chinese currency happened during a period of enormous international pressures on the Chinese government to appreciate its undervalued currency, the timing of the change is widely considered as "unexpected". There is much anecdotal evidence as well as academic studies supporting this statement. First, foreign pressures on the Renminbi for an appreciation had existed for more than two years, and the Chinese government regarded its exchange rate policy as a matter of China's sovereignty and rejected any political pressure on this issue. For example, on June 26, 2005, China's Premier Wen Jiabao said at the Sixth Asia-Europe Finance Ministers Meeting in Tianjin that China would "independently determine the modality, timing and content of reforms" and rejected foreign pressures for an immediate shift in the nation's currency regime. [9] One day later, Zhou Xiaochuan, the governor of the PBOC, said that it was too soon to drop the decade-old fixed exchange rate regime and that he had no plans to discuss the currency issue at the weekend meeting of the global central bankers in Basel, Switzerland. [10] On July 15, one week before the exchange

[9] See "Chinese premier warns against yuan reform haste" by the Wall Street Journal (http://online.wsj.com/article/0,,SB111975074805069620,00.html) Access date: October 9 2012

[10] See "China's Zhou Says `Time Is Not Ripe' to Drop Yuan Peg to Dollar" by Bloomberg (http://www.bloomberg.com/apps/news?pid=newsarchive&sid=a7n6HBTVapBA&refer=home) Access date: October 9 2012

rate system reform, the PBOC denied that it was planning to announce a revaluation of its currency. [11] On July 19, even two days before the reform, the PBOC still insisted that it would continue to keep the exchange rate stable and at a reasonable and balanced level in the second half of the year. [12]

Second, as elaborated by Yuan (2012), there was division in Chinese policy makers regarding whether the Chinese currency should be appreciated during that period. Specifically, the Ministry of Commerce opposed the currency appreciation (so as to maintain the competitiveness of China's export sector), while the other three central governmental agencies, the People's Bank of China, the National Development and Reform Commission, and the Ministry of Finance, all proposed revaluing the Chinese currency.

Third, after the reform, both the domestic and international media responded to the revaluation as completely unexpected. For example, CNN reported the episode as "The surprise move by China, ...". [13] The *Financial Times* wrote in its famous "Lex Column" on July 22, 2005 that "China likes to do things [in] its own way. After resisting pressure to revalue the Renminbi for so long, Beijing has moved sooner than even John Snow, the U.S. Treasury secretary, expected".[14] On July 22, 2005 the BBC *Worldwide Monitoring* said that "The People's Bank of China unexpectedly announced last night that the

[11] See "Central bank denies revaluation in August" by People's Daily (http://english.peopledaily.com.cn/200507/17/eng20050717_196621.html) Access date: October 9 2012

[12] See "China to keep RMB exchange rate basically stable: central bank" by People's Daily (http://english.peopledaily.com.cn/200507/20/eng20050720_197148.html) Access date: October 9 2012

[13] See "World events rattle futures" by CNN (http://money.cnn.com/2005/07/21/markets/stockswatch/index.htm) Access date: October 9 2012

[14] See "Renminimal THE LEX COLUMN" by Financial Times (http://www.lexisnexis.com.libproxy1.nus.edu.sg/ap/academic/) Access date: October 9 2012

Renminbi will appreciate by 2 per cent and will no longer be pegged to the US dollar".[15]

Fourth, academic studies also imply that the change in the exchange rate policy in July 2005 is unexpected. For example, Eichengreen and Tong (2011) study the impact of the Renminbi revaluation announcement on firm value in the 2005–2010 period. Using the change of stock prices before and after the announcement of the revaluation for 6,050 firms in 44 countries, they find that the Renminbi appreciation significantly increased firm values for those exporting to China while significantly decreased firm values for those competing with Chinese firms in their home markets, suggesting the exogeneity of the policy change.

1.2.3 Estimation Specification

The benchmark model (or its variants) used in the literature to investigate the response of exports to exchange rate is[16]

$$\ln V_{it} = \beta \ln e_{it} + \gamma_i + \eta_t + \varepsilon_{it}, \tag{1.1}$$

where V_{it} is the export value from Home country to foreign country i at time t; e_{it} is the nominal exchange rate of foreign country i's currency against the Home currency at time t; γ_i and η_t are the foreign country and time fixed effects, respectively; and ε_{it} is the error term.

[15] See "Hong Kong daily says exchange rate reform advantageous overall" by BBC Worldwide Monitoring (http://www.lexisnexis.com.libproxy1.nus.edu.sg/ap/academic/) Access date: October 9 2012

[16] For example, Kenen and Rodrik (1986) and Perée and Steinherr (1989) use a time series version of Equation (1.1) and find that the estimated coefficient β is smaller than 1 in most of their sample countries. Colacelli (2009) uses the same specification in a sample of 136 countries for the 1981-1997 period and also find a very small estimated coefficient β (equal to 0.055).

However, a crucial assumption to obtain an unbiased estimate of β in Equation (1.1) is that conditional on all the control variables, exchange rate is uncorrelated with the error term, i.e.,

$$E[\ln e_{it} \cdot \varepsilon_{it} \mid \gamma_i, \eta_t] = 0. \tag{1.2}$$

It is reasonable to doubt that this identifying assumption holds. For example, Dekle, Jeong and Ryoo (2010) show that producer heterogeneity is an important missing variable in the estimation of Equation (1.1). Meanwhile, export transactions involve buying and selling currencies, which in aggregate may influence the determination of the exchange rate. The violation of the identifying assumption (1.2) (due to the omitted variables bias and reverse causality) may explain why the literature only uncovered small values of β, which should theoretically be bigger than 1.[17]

To improve the identification, we use monthly data instead of the commonly-used yearly data, which precludes any potential omitted variables that do not vary monthly. Secondly, and more importantly, we use the sudden and unexpected exchange rate reform in China in July 2005 to conduct a difference-in-differences estimation. Specifically, we compare exports to the U.S. before and after July 2005 with exports to other countries during the same period. The DID estimation specification is:

$$\ln V_{it} = \delta Treatment_i \times Post_t + \gamma_i + \eta_t + \upsilon_{it}, \tag{1.3}$$

where $Treatment_i$ is the treatment status indicator, which takes value 1 if the country is the U.S. (the treatment group) and 0 otherwise (the control group); and $Post_t$ is the post-appreciation period indicator, which takes value 1 if it

[17] See Berman, Martin, and Mayer (2012) for the proof.

is after July 2005 and 0 otherwise. To adjust the potential serial correlation and heteroskedasticity, we use the robust standard error clustered at the country level (see Bertrand, Duflo, and Mullainathan, 2004).

Note that in the pre-revaluation period exchange rates of the Chinese currency (RMB) against other non-U.S. countries (say for example, the UK pound) were set by the cross rate between the dollar-RMB and the pound-dollar rate. If such approach was still applied and the pound-dollar rate did not change much in the post-revaluation period, the change in the pound-RMB exchange rate was then entirely driven by the change in the dollar-RMB rate, making our DID estimation strategy invalid. Two pieces of evidence help us relieve such concern. Firstly, after the revaluation in July 2005, the RMB was traded flexibly with a *reference basket* of currencies with the rate set by the PBOC every day; in other words, the cross-rate approach was largely not applied in the post-revaluation period. Secondly, as shown in Figure 1.2, the trade-weighted exchange rate of the RMB against other currencies remained quite stable in a two-year window of the revaluation time, and similar results are found for individual countries (results available upon request).

The identifying assumption associated with the DID estimation specification (1.3) is that conditional on a whole list of controls (γ_i, η_t), our regressor of interest, $Treatment_i \times Post_t$, is uncorrelated with the error term, υ_{it}, i.e.,

$$E\left[Treatment_i \times Post_t \cdot \upsilon_{it} \mid \gamma_i, \eta_t\right] = 0. \qquad (1.4)$$

As discussed in Section 1.3, the revaluation of the Chinese currency against the US dollar in July 2005 was highly unexpected, and therefore can be considered largely as an exogenous shock to Chinese exporters, which implies the satisfaction of the identifying assumption (1.4). Nonetheless, we conduct a battery of robustness checks to corroborate the claim that the identifying assumption (1.4) holds. These include a control for country-specific month effects and a country-specific linear time trend, a check on the pre-treatment differential trends between the treatment and control groups, a placebo test using homogeneous goods as the regression sample, and a difference-in-difference-in-differences (triple difference) estimation. For details, see Section 1.5.3.

1.3 Empirical Findings

1.3.1 Graphical Presentation

We start with a visual examination of the difference between Chinese exports to the treatment group (i.e., the U.S.) and the control group (i.e., other 55 countries) over time in Figure 1.3. The solid vertical line marks the time of China's exchange rate reform (i.e., July 2005), while the dashed vertical line represents one year before the reform. Arguably, the U.S. vs. non-U.S. export differential exhibits a four-stage pattern over our sample period (i.e., 2000-2006): from 2000 to late 2001, the export differential was quite stable; then it started a clear downward trend until the decline flattened out around mid-2004, or one year before the exchange rate reform in July 2005; and finally after the reform, Chinese exports to the U.S. decreased sharply against Chinese exports to the rest of our sample countries.

The above export-differential pattern coincides with that of the exchange rate differential displayed in Figure 1.2. For example, other currencies started to appreciate against the Chinese currency in early 2002 and stabilized around early 2004, during which period Chinese currency remained pegged to the US dollar. Between 2004 and 2006, while these other currencies stayed quite stable against the Chinese currency (despite of some ups and downs), US dollar began to continuously depreciate against Chinese currency after China's exchange rate reform in July 2005.

A few results emerge from these two figures. First, a currency appreciation has a visible, negative effect on exports as demonstrated by the negative correlation between the U.S. vs. non-U.S. export differential and their currency differential. Second, there is no clear differential pattern between U.S. and non-U.S. exports one year before the exchange rate reform, indicating that the reform is plausibly exogenous to exporters. Third, while after the reform in July 2005, the US dollar started to continuously depreciate against Chinese currency, other currencies remained quite stable throughout the period of 2004–2006, which justifies the use of the difference-in-differences estimation. However, as we include all sample periods in our analysis, one may be concerned that the results from the comparison of U.S. exports before and after the exchange rate reform with non-U.S. exports during the same period could be driven by the negative correlation between exports and currency changes happened during the period of 2002-2004. To address this concern, in a robustness check, we restrict our analysis to the period of 2004-2006.

1.3.2 Main Results

Regression results corresponding to Equation (1.3) are reported in Column 1 of Table 1.2. It is found that $Treatment_i \times Post_t$ is negative and statistically significant, implying that the appreciation of the Chinese currency against US dollar significantly reduced Chinese exports to the U.S. Meanwhile, the fall in exports is found to be substantial, i.e., the reform caused Chinese exports to the U.S. to fall by 17.6%.

In Column 2 of Table 1.2, we include monthly imports (in logarithm form), as the reform may make imports to China cheaper, and hence affect the production and exporting behavior of Chinese exporters (i.e., through the use of imported intermediate inputs and the increased domestic competition by imported final goods; see Amiti, Itskhoki, and Konings, 2014 for an elaboration on this point). In Column 3 of Table 1.2, we further include a measure of producer heterogeneity (i.e., the mean of export value divided by its standard deviation), the omission of which has been argued to seriously bias previous estimates in the literature (see Dekle, Jeong and Ryoo, 2010). Clearly, we find a quite similar negative estimate with the inclusion of these two additional controls.

Despite the fact that the reform was exogenous to Chinese exporters, one may be concerned that the decision to appreciate the currency in July 2005 by the Chinese central government was strategic. In other words, the drop in exports to the U.S. following the currency revaluation in July 2005 could have been driven by the U.S.-specific month effect, specifically, U.S.-July effect. To address such concern, we further include the country-specific month-of-year effect (i.e., $\gamma_i \times M_t$, where M_t is a month indicator such as

January, February, …, December), and the identification for example comes now from the comparison of U.S.-vs.-non-U.S. in July 2005 with U.S.-vs.-non-U.S. in July 2004. As shown in column 4 of Table 1.2, our main results regarding the effect of exchange rate on exports barely change in either statistical significance or magnitude, suggesting that our results are not driven by the country-specific month effect.

1.3.3 Robustness Checks

In this sub-section, we present a battery of robustness checks on our aforementioned estimation results.

Control for country-specific linear time trend. One concern is that it seems other currencies also started a depreciation trend against the Chinese currency after January 2005, continuing even after July 2005, the time of the exchange rate reform. To address the concern that our estimates may be contaminated by these similar depreciation time trends, we saturate the model with the inclusion of country-specific linear time trend, $\gamma_i \times t$. Hence, our identification comes from the discontinuity in the time trend caused by the revaluation of the Chinese currency against the US dollar in July 2005, a strategy similar to the regression discontinuity method. Despite of a significant drop in its magnitude, $Treatment_i \times Post_t$ remains negative and statistically significant (Column 1 of Table 1.3).

Check on pre-reform differential trends. A corollary of the identifying assumption (1.4) is that exports to the U.S. and other countries followed similar patterns before the revaluation in July 2005. Figure 1.3 clearly shows that U.S. vs. non-U.S. export differential was quite stable one year before the reform, but sharply declined right after the reform. To establish these results more formally, we first divide the whole 2000-2006 period into four periods (i.e., before July 2004, July 2004–June 2005, July 2005, and August 2005 onward), and then construct interactions between $Treatment_i$ and indicators of the three periods with July 2005 being the omitted category. The regression results are reported in Column 2 of Table 1.3. Consistent with the findings in Figure 1.3, the coefficient of $Treatment_i \times 07/2004-06/2005$ is highly insignificant, further confirming that U.S. exports and non-U.S. exports had similar patterns one year before the reform. Meanwhile, $Treatment_i \times Before\ 07/2004$ is positive and statistically significant, consistent with the fact that during this period there was a depreciation of the Chinese currency against other non-U.S. countries in our regression sample. Finally, our main results, the coefficient of $Treatment_i \times 08/2005\ onward$, remains negative and statistically significant. [18]

A sub-sample of the 2004-2006 period. As discussed in the Section 1.5.1, there is a concern that our findings of the negative impact of exchange rate appreciation on exports could be driven by the movement in earlier months, i.e., 2002-2004. Meanwhile, the exchange rate of currencies other than the US dollar remained quite stable against the Chinese currency during the period of 2004-2006, making the difference-in-differences analysis using

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[18] Interestingly, the coefficients of $Treatment_i \times 08/2005\ onward$ and $Treatment_i \times Before\,07/2004$ have similar magnitudes but opposite signs, as the former captures the appreciation effect of the dollar-RMB rate with other exchange rates unchanged , while the latter reflects the depreciation effect of the Chinese currency against other non-U.S. countries with the dollar-RMB rate unchanged.

just the data of 2004-2006 more appealing. To these ends, we conduct a robustness check by restricting our analysis to the sample of 2004-2006. Regression results are reported in Column 3 of Table 1.3. Despite of a drop in the estimated magnitude, $Treatment_i \times Post_t$ remains negative and statistically significant, implying the robustness of our previous findings.

A placebo test using homogeneous goods. The identification from our difference-in-differences estimation comes from the fact that the exported goods are priced differently across the treatment and control groups, and hence the appreciation of the treatment country's currency makes the exported goods more expensive in the treatment country, thus producing a fall in total exports to that country, given that the situations in the control group remain unchanged. However, if the exported goods are charged with the same prices across countries and hence the export prices are detached from the exchange rate, then we should not spot any significant effects from the difference-in-differences estimation. One example of these special exported goods are commodities traded on the exchange market, or the group of homogeneous goods as classified by Rauch (1999). Using Rauch (1999)'s classification, we divide the whole set of Chinese exported goods into two groups, differentiated and homogeneous goods, and then conduct a placebo test using the sample of homogeneous goods. The regression results are reported in Column 4 of Table 1.3. Consistent with our argument, the coefficient of $Treatment_i \times Post_t$ is highly insignificant, lending further support to our identification.

A difference-in-difference-in-differences estimation. Further exploring the difference between differentiated and homogeneous goods, we conduct a difference-in-difference-in-differences (or triple difference) estimation. Specifically, we estimate the following equation:

$$\ln V_{igt} = \delta Treatment_i \times Post_t \times Differentiated_g + X'_{igt}\varphi + \gamma_{ig} + \eta_{gt} + \chi_{it} + \upsilon_{igt}, \quad (1.5)$$

where g indicates the group of the exported goods, i.e., differentiated or homogeneous goods group; $Differentiated_g$ is an indicator of the differentiated goods group; and X_{igt} is a vector of controls (i.e., the logarithm of imports and producer heterogeneity).[19]

The beauty of the triple difference estimation is that it allows us to include a full set of the country-group fixed effects γ_{ig}, the group-time fixed effects η_{gt}, and the country-time fixed effects χ_{it}.[20] For example, the inclusion of the country-month fixed effects means controlling for all observed or unobserved time-invariant and time varying country characteristics, which are the main concerns violating our above difference-in-differences identifying assumption (1.4). As shown in Column 5 of Table 1.3, the triple interaction term is found to be negative and statistically significant. This further reinforces our aforementioned difference-in-differences estimation results, i.e., our findings are not biased due to some omitted time-varying country characteristics.

---

[19] Note that the number of observations increase as the regression unit is now at the group-country-month level.

[20] In estimating the Equation(1.5), we first difference exports across the two groups within a country-month cell, and then estimate the resulting double-difference equation:

$$\ln \tilde{V}_{it} = \delta Treatment_i \times Post_t + \tilde{X}_{it}\varphi + \gamma_i + \eta_t + \tilde{\upsilon}_{it},$$

where tilded variables mean cross-group differenced, e.g., $\ln \tilde{V}_{it} \equiv \Delta \ln V_{igt}$. Otherwise, we encounter the computational burdens as the original triple difference equation involves too many dummy variables, i.e., $56 \times 84 = 4{,}704$ country-time dummies, $56 \times 2 = 112$ country-group dummies and $84 \times 2 = 168$ group-time dummies.

1.3.4 Exchange Rate Elasticity

Although in the previous sections we have established that the exchange rate reform (or the currency appreciation) has a negative effect on exports, it is interesting to know the exchange rate elasticity of exports. To this end, we use the exchange rate reform in China to construct an instrumental variable for the exchange rate and estimate Equation (1.1) with the two-stage-least-squares (2SLS) method.[21]

We start with the estimation of Equation (1.1) without instrumenting the exchange rate in column 1 of Table 1.4. Though statistically significant, the estimated coefficient of exchange rate has only a value of -0.454, a magnitude similar to those found in the literature (e.g., Colacelli, 2009).

The instrumental variable estimation results are reported in Column 3 of Table 1.4. The first-stage results (not included here but available upon request) show a positive and statistical relation between the instrument ($Treatment_i \times Post_t$) and the regressor of interest ($\ln ER_{ct}$). And the F-test of excluded instruments in the first-stage has a value of 27.02, substantially higher than the critical value 10 of the "safety zone" for strong instruments suggested by Straight and Stock (1997). These results suggest that our proposed instrument is both relevant and strong.

With respect to our central issue, the exchange rate, after being instrumented, still casts a negative and statistically significant impact on total exports. More importantly, there is a substantial increase in the estimated

[21] For the exclusion restriction to be hold, we need that the exchange rate reform affects exports value only through its effect on the bilateral exchange rate, and not through other changes in the economy that might be coincident with the exchange rate reform.

magnitude: a 1% appreciation causes total exports to fall by 1.61%, confirming the theoretical prediction that the exchange rate elasticity of exports is greater than 1 and the existence of a significant bias in the previous OLS estimations. Put the number into a real context: given that China exported US$1.904 trillion worth of goods in 2011, a 1% currency appreciation means a US$30.65 billion loss in China's export sector, a significant number justifying why government officials and businessmen are greatly concerned about the currency appreciation.

In Column 2 and 4 of Table 1.4, we replace the nominal exchange rate with the real exchange rate. Clearly, we still identify a statistically significant effect of exchange rate on total exports, although the magnitude of the IV estimate drops from -1.605 to -1.125.

1.3.5 Trade Diversion

From a policy viewpoint, it is important to know whether the fall in exports to the treatment group (i.e., the U.S.) after the currency appreciation causes a withdrawn by Chinese exporters from the exporting market or the diversion from the affected destination (i.e., the U.S.) to some unaffected destinations in our regression sample. If it is the latter, then for governments, the prospects after a currency appreciation may not be that gloomy.

Based on the premise that it is easier to divert exports to countries (such as other OECD countries) with similar consumer preference as those of the U.S., we conduct two exercises to shed light on the possibility of trade diversion. Firstly, we exclude OECD countries from our control group and re-estimate Equation (1.3). If there were trade diversion, we should expect a smaller estimation coefficient. However, we find in Column 1 of Table 1.5 that the coefficient of $Treatment_i \times Post_t$ slightly to -0.186 from -0.165 (in Column 4 of Table 1.2; with all countries in the regression), despite of the increase being statistically insignificant.

Secondly, we compare Chinese exports to OECD countries (excluding the U.S.) before and after the exchange rate reform with the corresponding exports to the rest of countries in our sample during the same period. If there were trade diversion, we should expect that following the appreciation of Chinese currency against the US dollar, Chinese exports to other OECD countries have increased relative to Chinese exports to other sample countries, given that these countries' currencies remained stable against Chinese currency during this period. However, as shown in Column 2 of Table 1.5, $Treatment_i \times Post_t$ is highly insignificant.

These two exercises demonstrate that there is no substantial evidence to support trade diversion hypothesis after the exchange rate reform, and much of the falls in Chinese exports to the U.S. are due to the exits of Chinese exporters from the exporting market.

1.3.6 Mechanism

While our objective is to investigate the export response to changes in the exchange rate at the macro-level, our customs data contain observations disaggregated at the firm-product-month-country level, which allows us to investigate some underlying mechanism about how currency appreciation affects total exports. In the Appendix 1, we show that the effect of exchange rate changes on aggregate exports operates on two margins, the intensive- and the extensive-margins. Specifically, a currency appreciation causes the final price in the foreign market to increase and the FOB export price to decrease, due to an incomplete pass-through. The final price increase may reduce the demand, which, combined with the decreased FOB price, will reduce the total export revenue, damping the effect of the appreciation at the intensive margin. Moreover, the adverse effect of a currency appreciation is stronger for less productive exporters, making them unprofitable in and hence exit the foreign market (an extensive-margin effect).

The regression results are reported in Table 1.6. In Column 1-2, we investigate the extensive-margin effect, that is, regressing the total number of firms and the total number of HS-8 product categories exported to the U.S. on $Treatment_i \times Post_t$ along with a full set of controls. It is found that, consistent with our model featuring heterogeneous firms, the Chinese currency appreciation significantly reduced the number of total exporters and the number of HS-8 product categories, specifically, by 6.6% and 29.2%, respectively, in magnitude.

In column 3-5, we investigate the intensive-margin effect from different dimensions as suggested by the model. Specifically, we focus on the sample of surviving exporters (firms continuing to export after the currency appreciation) and regress the mean values of export price, export volume and export revenue at the firm-product-month-country level on $Treatment_i \times Post_t$ along with a full set of controls. Our model predicts that, due to incomplete pass-through, the appreciation of Renminbi will decrease the FOB export price. This prediction is confirmed by the estimate in Column 3, i.e., the appreciation brings down the price by about 1.3%, which is very significant both statistically and economically. Also consistent with the model, the effect on export volume (shown in Column 4) is found to be negative, albeit not precisely estimated. The total intensive margin effect of the Renminbi appreciation is shown in column 5. Given the negative effects of the appreciation on the price and the volume, it is natural that the appreciation has strong negative impact on export revenue, i.e., a fall of 4.1%.

In summary, we find support for both extensive-margin and intensive-margin effects of exchange rate movement on exports.

1.3.7 Heterogeneous Effects

In the last part of our empirical investigation, we examine possible heterogeneous effects across different regions (i.e., inland versus coastal regions), across different types of firms (i.e., state-owned enterprises versus private enterprises), and across different industries/products (i.e., time sensitive versus time insensitive industries; different product categories in the PPI basket). The estimation specification we use is the triple difference Equation (1.5), with different definitions of the group indicator in different investigations.

Coastal versus inland regions. We start in Column 1 of Table 1.7 the investigation of differential exports response to exchange rate changes between coastal and inland regions. The group indicator takes a value of 1 for coastal regions and 0 for inland regions. The triple interaction term is found to be negative and statistically significant, indicating that exports to U.S. fell more in coastal regions than in inland regions after the appreciation of the Chinese currency against the US dollar. One possible explanation is that as the transport costs are lower in coastal regions, the initial cut-off productivity levels of exporting is lower in coastal regions than in inland regions. The currency appreciation increases the cut-off productivity levels of exporting in both coastal and inland regions, but as there are much weaker exporters in coastal regions, more exporters from coastal regions exit the exporting market than their counterparts from inland regions. [22]

State-owned versus private enterprises. In Column 2 of Table 1.7, we investigate the possible different responses between state-owned enterprises and private enterprises, with the group variable indicating a state-owned enterprise. Clearly, we find that state-owned enterprises respond more to exchange rate changes than private enterprises, i.e., the former's exports fall more than the latter's. One possible explanation is that state-owned enterprises in China receive many subsidies from the governments (such as trade credit, export rebate, etc), making the cut-off productivity levels of exporting for state-owned enterprises lower than those for private enterprises. Then after the currency appreciation, some weaker state-owned enterprises are driven out of the exporting market, if the government subsidies remain rigid in the short-run.

Time sensitive versus time insensitive industries. Thirdly, we divide industries into two groups, time sensitive (assigned a value of 1 for the group indicator) and time insensitive industries (assigned a value of 0 for the group indicator), following the classification used by Djankov, Freund, and Pham (2013). Specifically, time sensitive industries are the three 2-digit manufacturing industries (i.e., office equipment, electric power machinery,

[22] Using the theoretical model laid out in the Appendix 1, we can show that $\frac{\partial^2 V_i}{\partial e_i \partial \tau_i} > 0$.

and photographic equipment) having the highest probability of using air transport, whereas time insensitive industries are the three 2-digit manufacturing industries (i.e., textile yarns, cement, and plumbing fixtures) with the lowest probability (the probability was estimated by Hummels, 2001). As shown in Column 3 of Table 1.7, time sensitive industries experienced more of a fall in exports after the revaluation of exchange rate in July 2005 than time insensitive industries. One possible explanation is that production and shipment are easier to adjust and hence more responsive to exchange rate movements in time sensitive industries than in time insensitive industries.

Different product categories in the PPI basket. Finally, following Vermeulen et al. (2007), we group products into 6 categories used in the PPI basket (i.e., food products, non-food non-durable products, durable products, intermediate goods, energy, and capital goods) to examine whether there are differential appreciation effects. Regression results are reported in columns 4-9 of Table 1.7. It is found that the currency appreciation had significant and negative effects on exports in non-food non-durable products, durable products, intermediate goods and capital goods, but insignificant effects in food products and energy. Intuitively, food and energy are necessities of life, which make them non-responsive to exchange rate changes.

1.4 Conclusion

The effect of exchange rate changes on exports has attracted extensive attention from policy makers, commercial circles and academia. In this paper, we revisit the question of whether exports respond to exchange rate changes and contribute to the literature by carefully addressing the identification issues.

Specifically, we employ monthly, rather than yearly data usually used in the literature, to take advantage of more variations in the key variables. And to address the potential endogeneity problem in the estimation, we use the unexpected exchange rate regime switch by Chinese government in July 2005 as a natural experiment.

The difference-in-differences estimation uncovers a statistically and economically significant and negative effect of a currency appreciation on exports. Specifically, our main estimation result shows that a 1% exchange rate appreciation decreases total exports by 1.61%, which, in the context of year 2011 China, represents a US$30.65 billion decrease in total exports. This negative effect is robust to various checks on the validity of the difference-in-differences estimation and other econometric concerns. Meanwhile, we do not find any trade diversion by Chinese exporters after the currency appreciation, but uncover both intensive-margin and extensive-margin effects of exchange rate changes on exports, and heterogeneous effects across regions, firms, and industries/products.

Chapter Two

Exchange Rates and Export Structure

2.1 Introduction

Exchange rates have been an important tool of trade policies. A weaker currency is widely believed by politicians and government officials to stifle import competition, helping to relieve domestic political pressures from high unemployment rates and boosting the performance of export sectors, subsequently leading to economic growth. Substantially hit by the 2008-09 financial crisis, developed economies like the U.S., Japan, and European countries have altered their monetary policies, which has deliberately or unintentionally caused their currencies to depreciate. Many developing countries also purposely undervalue their currencies by a fixed-exchange-rate regime or constant interventions to pursue a long-run export-led growth strategy.[23] International politics hence often involves the scenario where the developed countries ask the developing ones to appreciate their currencies.

Nevertheless, firms and industries respond to exchange-rate movement differently. For example, Berman, Martin, and Mayer (2012) find that by reducing their markups, more productive exporters can absorb negative shocks of currency appreciation better than their less productive counterparts. At the sectoral level, if appreciation of a developing country's currency moves its export structure towards the industries in which developed countries are concentrated in, the corresponding depreciation of developed countries'

[23] In the case of China, reliable estimates show that Chinese currency was undervalued by around 40% as of 2000 (Frankel 2006) and around 25% as of 2005 (Rodrik 2010). Rodrik (2008) explains this rationale by showing the clear positive associations between undervalued currencies, large exports, and rapid growth in developing countries.

currencies may thus have limited effect on restraining imports and promoting exports.

To the best of our knowledge, there is no work on how the exchange rate changes a country's export structure (i.e., the distribution of export values across different industries), despite numerous studies on the effect of the exchange rate on aggregate export values and individual firm behaviors (e.g., Amiti, Itskhoki, and Konings, 2014; Berman, Martin, and Mayer, 2012; Chatterjee, Dix-Carneiro, and Vichyanond, 2013; Dekle, Jeong, and Ryoo, 2010; Li, Ma, and Xu, 2013). This paper fills this void by using a sudden and unexpected currency revaluation in China to examine whether and how the exchange rate affects export structure.

On July 21, 2005, the Chinese government unexpectedly revalued its currency against the U.S. dollar, which resulted in an immediate appreciation of 2.1 percent (for a detailed description on this episode and the unexpectedness, see Section 2.3). The sharp change in China's exchange rate provides us an opportunity to have an arguably clean identification of the effect of currency appreciation using a regression discontinuity (RD) estimation. Specifically, the exogeneity of currency appreciation makes export structure before currency appreciation (i.e., January 2005-July 2005) a good counterfactual to the one after currency appreciation (i.e., August 2005-December 2005). Meanwhile, to purge the monthly effect (e.g., differences in U.S. demand across months), we add data of a year during which Chinese currency was fixed against the U.S. dollar, as a control group, and conduct a difference-in-differences (DD) estimation.

In our empirical investigation, we use an index developed by Hausmann, Hwang, and Rodrik (2007) which measures how relatively heavily a good is exported by developed countries. In particular, we use this index to construct an export similarity index that measures how similar China's exports are to developed countries (see details in Section 2.3). Our RD-DD estimation results show that after the currency appreciation, China's export structure to the U.S. becomes more similar to that of developed countries.[24] These results remain robust to a battery of sensitivity checks, including a difference-in-difference-in-differences (DDD) estimation, a placebo test, an examination of U.S. exports to China, and an exclusion of processing trade.

To illustrate how the exchange rate changes export structure, we present a trade model with monopolistic competition in which two sectors of differentiated goods differ mainly in their elasticities of substitution. As the Chinese currency is heavily controlled and undervalued, we take the fact of an undervalued South's currency as the key feature defining the North-South structure. As explained in Section 2.4, there is strong evidence that developed countries export relatively heavily in goods with low elasticity of substitution (high markups). Given that the North exports relatively heavily in goods with high markups, we show that if the South's currency appreciates, the South's export structure becomes closer to the North's. The intuition is that when the South's exports become more expensive due to currency appreciation, the reductions in the North's expenditure on these goods are larger in the sector with higher price elasticity. Whereas this argument based on the intensive

[24] We focus on China's exports to the U.S. as China's sudden exchange-rate change is against the U.S. dollar. We do not examine China's exports to the world because the weighted average of exchange rates against various countries was quite volatile in 2005, as well as in other years.

margin with entry fixed in the short run fits our empirical results, the same result holds in the long run when free entry is allowed.

One direct implication of our empirical results is that since developed countries (or the U.S.) concentrate on and export relatively more of those goods with low elasticity of substitution, the competition in these goods from China is reduced, but not by much. Whereas our empirical results are necessarily short-term by the identification strategy, changes in export structure may have important long-run implications, especially with the resource reallocation and learning-by-doing effects so that Chinese producers may gradually become more productive and provide fiercer competition in these, so to speak, high-end sectors. Another long-run implication is related to the result in Hausmann, Hwang, and Rodrik (2007) that when a country's export structure becomes similar to that of developed countries, the ensuing economic growth of the country would be higher. Similar empirical findings are uncovered by Jarreau and Poncet (2012) in the context of China. The rationale is based on a "cost discovery" story or more generally, the idea of "countries become what they produce". In sum, whereas consumers in the South obviously would benefit from the South's currency appreciation, appreciation may not be all that bad even from the viewpoint of production.

While the model displays a mechanism of resource reallocation across firms within a locality, our empirical estimates capture the whole spectrum of resource reallocation. That is, our estimates capture three margins of the changes in export structure: across cities, within city and across firms, and within firm and across products. Meanwhile, by further exploring the data, we can decompose the appreciation effect on export structure into these three margins. We find that resource reallocation within city and across firms accounts for the majority of our appreciation effect (i.e., 72.22 percent), while resource reallocation across cities as well as within firm and across products explain 16.67 percent and 11.11 percent, respectively.

The paper is organized as follows. Section 2.2 provides a theory of how export structure is affected by the exchange rate. Section 2.3 describes our data, variables, and empirical strategy, including details of the reform of China's exchange-rate regime in July 2005. Empirical results including robustness checks are reported in Section 2.4, and Section 2.5 concludes.

2.2 A Model of Exchange Rate and Export Structure

We extend a standard monopolistic competition model of trade à la Krugman (1980) and Helpman and Krugman (1985) to provide a plausible mechanism regarding how export structure is affected by the exchange rate.

2.2.1 Model Setup

There are two countries, North and South, with population L_n and L_s, respectively. Here, we think of China as the South, who sets up a fixed-exchange-rate regime, and therefore the exchange rate between two countries is a policy (exogenous) variable. There are three goods/industries in the economy, and the utility of a representative agent in country j follows a Cobb-Douglas form: $U_j = Q_{j0}^{\alpha 0} Q_{j1}^{\alpha 1} Q_{j2}^{\alpha 2}$, where $\alpha_i \in (0,1)$, for $i \in \{0,1,2\}$, Q_{ji} is the consumption of good i, in country j, and $\sum_i \alpha_i = 1$. Labor is the only production input. Good 0 is the numeraire good produced with a constant returns technology and is freely traded within and between countries. This numeraire good is not subject to currency exchange. We normalize the labor productivity of good 0 to 1, and hence wages are also normalized to 1 in both countries.

Goods 1 and 2 are both differentiated and tradable, and the composite Q_{ji}, is made by

$$Q_{ji} = \left(\int_{\Omega_{ji}} [q_{ji}(\omega)]^{\frac{\sigma_i - 1}{\sigma_i}} d\omega \right)^{\frac{\sigma_i}{\sigma_i - 1}},$$

where $q_{ji}(\omega)$ is the consumption of variety ω, and Ω_{ji}, denotes the set of the varieties of good i, consumed in country j. The elasticity of substitution is σ_i in industry i. We assume that $\sigma_2 < \sigma_1$, so that good 2 has a lower price elasticity than good 1. Trade in the two differentiated industries is subject to currencies and the exchange rate, i.e., people sell and buy the goods with the country's currency if the trade is within the country, and if trade is between countries, then currency exchange is needed. Barring frictions, the real exchange rate of these goods across countries is 1. However, there are numerous factors/distortions that will create a bias of the real exchange rate from 1. Especially in the fixed-exchange-rate regime, the real exchange rate may differ significantly from 1. Say, a unit of a good in the U.S. can be exchanged for $e < 1$ units of the same good in China (hence one unit of good in China can be exchanged for $e^{-1} > 1$ units in the U.S.). From here onward, we assume that the real exchange rate from a North's to a South's good is $e<1$, which captures the fact that the South often uses the exchange rate as a policy tool to implement an export-oriented development strategy.

On top of the exchange-rate distortion, trade between countries is also subject to standard iceberg trade cost so that to deliver one unit to the other country, $\tau>1$ units needs to be shipped. By paying an entry cost κ, each firm draws a distinct variety (and hence is a monopolist for it) and can produce the good with constant marginal cost c. Firms can price discriminate across countries. The probability that a variety will be in industry i is given by λ_i, and $\lambda_1+\lambda_2=1$. Free entry determines the number of firms M_j in each country j. The number of firms in industry i in country j is therefore $M_{ji}=\lambda_i M_j$.

Note a key difference between e and τ in the model.[25] Here, an increase of the trade cost τ increases import prices in both countries and the degree of separation between the two markets, whereas a decrease in e increases the South's import prices while decreasing the North's import prices. Hence, e has an asymmetric effect, whereas the effect of τ is symmetric. Having multiple sectors with different σ_i and the asymmetric effect of e considerably increases the complexity of the model, and hence for tractability and for our purpose of illustrating sectoral shifts, we opt to go with a homogeneous-firm model, instead of a heterogeneous-firm one.

[25] As will be more clear after Proposition 1, the role of trade cost τ is indispensible, because without it, i.e., $\tau=1$, there won't an equilibrium, since all the firms will earn more profit in the South than the North, making it impossible for the free entry condition to hold in both countries.

2.2.2 Equilibrium and the Effect of the Exchange Rate

Let $p_{ji}(\omega)$ be the price of variety ω of industry i that faces the consumers in j. The Cobb-Douglas-CES structure implies that the total sales of variety ω of industry i in country j is

$$r_{ji}(\omega) \equiv p_{ji}(\omega) q_{ji}(\omega) = \alpha_i L_j \left(\frac{p_{ji}(\omega)}{P_{ji}} \right)^{1-\sigma_i},$$

where P_{ji} is the standard price index $P_{ji} = \left(\int_{\Omega_{ji}} p_{ji}(\omega)^{1-\sigma_i} d\omega \right)^{\frac{1}{1-\sigma_i}}$. Let p_{ni}^I denote the price of an imported good in the North (from a South firm).[26] A South's firm profit is

$$\begin{aligned} \pi_{si} &= (p_{si} - c) q_{si}(p_{si}) + (e^{-1} p_{ni}^I - \tau c) q_{ni}(p_{ni}^I) \\ &= (p_{si} - c)(p_{si})^{-\sigma_i} \frac{\alpha_i L_s}{P_{si}^{1-\sigma_i}} + (e^{-1} p_{ni}^I - \tau c)(p_{ni}^I)^{-\sigma_i} \frac{\alpha_i L_n}{P_{ni}^{1-\sigma_i}}. \end{aligned}$$

Equilibrium pricing follows a standard markup rule, where the markup is denoted as $\mu_i = \dfrac{\sigma_i}{\sigma_i - 1}$. In particular, the effective (delivered) marginal cost τc is incurred in the South, and the price in the South's viewpoint is $e^{-1} p_{ni}^I = \mu_i \tau c$. Hence, $p_{si} = \mu_i c, p_{ni}^I = \mu_i e \tau c$, and the profit of a South's firm is

$$\pi_{si} = \alpha_i (\mu_i - 1) \mu_i^{-\sigma_i} c^{1-\sigma_i} \left(\frac{L_s}{P_{si}^{1-\sigma_i}} + e^{-\sigma_i} \tau^{1-\sigma_i} \frac{L_n}{P_{ni}^{1-\sigma_i}} \right).$$

Similarly, for the North, we have $p_{ni}(c) = \mu_i c, p_{si}^I(c) = \mu_i e^{-1} \tau c$, and π_{ni} is similarly derived. The price indices are rewritten as

---

[26] Note that p_{ni}^I is the sales of one unit of a good with the North currency (but denominated in numeraire), and these convert to more than enough South currency to buy one unit $(e^{-1} > 1)$.

$$
\begin{aligned}
P_{si}^{1-\sigma_i} &= \lambda_i (\mu_i c)^{1-\sigma_i} [M_s + M_n (e^{-1}\tau)^{1-\sigma_i}], \\
P_{ni}^{1-\sigma_i} &= \lambda_i (\mu_i c)^{1-\sigma_i} [M_n + M_s (e\tau)^{1-\sigma_i}]
\end{aligned}
$$

An entrant's expected profit in the South is

$$
\begin{aligned}
E\pi_s &= \lambda_1 \pi_{s1} + \lambda_2 \pi_{s2} - \kappa = \lambda_1 \alpha_1 (\mu_1 - 1)\mu_1^{-\sigma_1} c^{1-\sigma_1} \left(\frac{L_s}{P_{s1}^{1-\sigma_1}} + e^{-\sigma_1} \tau^{1-\sigma_1} \frac{L_n}{P_{n1}^{1-\sigma_1}} \right) \\
&+ \lambda_2 \alpha_2 (\mu_2 - 1)\mu_2^{-\sigma_2} c^{1-\sigma_2} \left(\frac{L_s}{P_{s2}^{1-\sigma_2}} + e^{-\sigma_2} \tau^{1-\sigma_2} \frac{L_n}{P_{n2}^{1-\sigma_2}} \right) - \kappa
\end{aligned}
$$

Zero expected profit condition is then $E\pi_s = 0 = E\pi_n$, which entails

$$
\frac{\alpha_1}{\sigma_1}\left[\frac{L_n(1-e^{-\sigma_1}\tau^{1-\sigma_1})}{m+(e\tau)^{1-\sigma_1}} + \frac{L_s(e^{\sigma_1}\tau^{1-\sigma_1}-1)}{1+m(e^{-1}\tau)^{1-\sigma_1}} \right] = \frac{\alpha_2}{\sigma_2}\left[\frac{L_n(e^{-\sigma_2}\tau^{1-\sigma_2}-1)}{m+(e\tau)^{1-\sigma_2}} + \frac{L_s(1-e^{\sigma_2}\tau^{1-\sigma_2})}{1+m(e^{-1}\tau)^{1-\sigma 2}} \right], \tag{2.1}
$$

where $m \equiv \frac{M_n}{M_s}$ is the ratio of entry between the two countries. The equilibrium entry ratio m^* satisfies (2.1), and the level of M_s and M_n can be determined by $E\pi_s = 0$ (or, equivalently, $E\pi_n = 0$). In the following proposition, we show that when trade cost τ is sufficiently large, there is a unique finite equilibrium entry ratio $m^* > 0$, which implies that equilibrium entries in both countries are positive. Moreover, m^* strictly increases with an appreciation of the South's currency.

Proposition 1 Denote $\ell = L_n / L_s$. Let τ_a be the solution of τ to the following equation.

$$
e^{\sigma}\left[\tau^{\sigma-1}\left(e^{-1}\ell\right) + \tau^{1-\sigma}\right] = e^{-1}\ell + 1,
$$

and

$$
\tau_b \equiv \begin{cases} \max\left\{1, 2^{\frac{1}{1-\sigma}}\left[e^{\sigma}\left(1+\ell e^{-1}\right) + \sqrt{e^{2\sigma}\left(1+\ell e^{-1}\right)^2 - 4\ell e^{-1}} \right]^{\frac{1}{\sigma-1}} \right\} & \text{if} \quad 4\ell e^{-1} \le e^{2\sigma}\left(1+\ell e^{-1}\right)^2 \\ 1 & \text{if} \quad 4\ell e^{-1} \le e^{2\sigma}\left(1+\ell e^{-1}\right)^2 \end{cases}
$$

Let $\hat{\tau}_i = \max\{\tau_{ai}, \tau_{bi}\}$, where τ_{ai} and τ_{bi} are the values of τ_a and τ_b when $\sigma = \sigma_i$. Suppose the trade cost τ is such that $\tau > \hat{\tau} \equiv \max\{\hat{\tau}_1, \hat{\tau}_2\}$. Then, there exists a unique finite equilibrium entry ratio $m^* > 0$ (positive entries in both countries), and m^* strictly increases in e.

Proof. See the Appendix 2

To understand this proposition, think of the case of $\tau = 1$ and $L_n = L_s$. In this case, there is no separation between the two countries, and the two countries are symmetric, except that the South's firms enjoy an edge due to exchange-rate distortion ($e < 1$). Hence, all firms in the South enjoy larger profits than those in the North, and $m^* = 0$ in equilibrium ($M_n = 0$). On the other hand, if $\tau \to \infty$, then the effect of $e < 1$ becomes nil and there must be positive entries in both countries. Hence, a sufficiently large τ is required to have enough separation between the two markets.[27] Since an increase in e implies that the South's firms' edge due to the exchange rate is reduced, and hence we expect less entry in the South and more in the North, leading to an increased m^*.

2.2.3 Export Structure and the Exchange Rate

Here, we first want to investigate the conditions under which $\frac{X_{s2}}{X_{s1} + X_{s2}} < \frac{X_{n2}}{X_{n1} + X_{n2}}$, that is, the more developed country's (North's) export

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[27] The condition also involves the ratio of country size $\ell = L_n / L_s$ because it is possible that given an e and τ, m^* becomes infinity ($M_s = 0$) when ℓ is very large so that the advantage of the South due to e is reversed due to the large population in the North and the home market effect. Nevertheless, regardless of the value of e and ℓ, as long as τ is sufficiently large, positive entries in both countries are guaranteed.

in industry 2 is more than that of the less developed country (the South). This is equivalent to $\frac{X_{s2}}{X_{s1}} < \frac{X_{n2}}{X_{n1}}$. We also want to investigate whether $\frac{d}{de}\frac{X_{s2}}{X_{s1}} > 0,$ and $\frac{d}{de}\frac{X_{n2}}{X_{n1}} < 0,$ so that the export structure of the two countries become more similar when the South's currency appreciates. Note that export volume from the South in industry $i,$ is $X_{si} = M_{si}e^{-1}p_{ni}^{I}q_{ni}\left(p_{ni}^{I}\right)$.

So,

$$\frac{X_{s2}}{X_{s1}} = \frac{\alpha_2\mu_2^{1-\sigma_2}\lambda_2}{\alpha_1\mu_1^{1-\sigma_1}\lambda_1}\frac{P_{n1}^{1-\sigma_1}}{P_{n2}^{1-\sigma_1}}\left(e\tau c\right)^{\sigma_1-\sigma_2}. \tag{2.2}$$

Similarly,

$$\frac{X_{n2}}{X_{n1}} = \frac{\alpha_2\mu_2^{1-\sigma_2}\lambda_2}{\alpha_1\mu_1^{1-\sigma_1}\lambda_1}\frac{P_{s1}^{1-\sigma_1}}{P_{s2}^{1-\sigma_1}}\left(e^{-1}\tau c\right)^{\sigma_1-\sigma_2}.$$

(2.3)

In the short run, M_n and M_s (and hence m) are fixed. If price indices were also fixed, then obviously $\frac{X_{s2}}{X_{s1}}$ increases with e, as $\sigma_1 > \sigma_2$. This is basically an intensive margin effect that when the South's goods become more expensive, the quantities demanded and sales in the North for these goods are reduced, but the effect is stronger for good 1 than good 2, because good 1 has a larger price elasticity. Proposition 2 shows that this effect at intensive margin is robust when taking into account the adjustment of price indices and free entry in the long run. It also provides two sufficient conditions under which $\frac{X_{n2}}{X_{n1}} > \frac{X_{s2}}{X_{s1}}$ holds, and hence the export structures in the two countries become more similar with a currency appreciation.

Proposition 2 Suppose that $\sigma_2 < \sigma_1$, $e \le 1$, and $\tau > \hat{\tau}$ so that there is a unique equilibrium with positive entries in both countries (**Proposition 1**). Then,

1. Both in the short run when entries M_n and M_s are fixed and in the long run when entries are determined by free entry, $\frac{d}{de}\left(\frac{X_{s2}}{X_{s1}}\right) > 0$ and $\frac{d}{de}\left(\frac{X_{n2}}{X_{n1}}\right) < 0$. That is, the South's export in industry 2 relative to that in industry 1 increases when currency in the South appreciates.
2. If one of the following conditions holds, then in equilibrium $\frac{X_{n2}}{X_{n1}} > \frac{X_{s2}}{X_{s1}}$, and the export structure in the South becomes closer to that in the North when the South's currency appreciates.

(a) The two countries have the same population size, i.e., $L_n = L_s$, and the real exchange rate is such that $e < 1$.

(b) The South has a larger population, i.e., $L_n > L_s$, and the real exchange rate is $e = 1$.

Proof. See the Appendix 2.

Given the empirical finding in the next subsection that developed countries export relatively more goods with low elasticity of substitution, the more important message of Proposition 2 is Point 1, because given this fact, currency appreciation leads to a more similar export structure. Point 2 shows some conditions under which the above-mentioned fact can be generated from the model. The intuition behind Point 2(a) is that $e<1$ creates an advantage for producers in the South, and this advantage is more pronounced for industry 1 because the price elasticity is larger. Although we do not model how the wages are determined, it is worthwhile noting that the price advantage of the South reflected by $e<1$ is similar to the effect when the South's wages are lower than the North's, which is fitting to the U.S.-China scenario. Point 2(b) holds mainly because the home market effect is more pronounced for the good with larger price elasticity. It is easy to verify numerically that the same result holds in the convex combination of these two conditions, i.e., the case of $L_s \geq L_n$ and $e \leq 1$.[28]

2.2.4 Developed Countries Export Relatively More Goods with Low Demand Elasticities

Our theoretical analysis shows that when the South appreciates its currency, its exports become more skewed towards the industry with lower elasticity of substitution, and the export structure becomes more similar to developed countries. To connect our theoretical and empirical analyses, it is important to examine whether developed countries export relatively more goods with low elasticity of substitution. To this end, we examine the correlation between two

[28] It is also possible to explain the difference in export structure via technological differences. One can think of this as $\lambda_n = \frac{\lambda_{n2}}{\lambda_{n1}} > \frac{\lambda_{s2}}{\lambda_{s1}} = \lambda_s$, i.e., the North firm is more able and hence more likely to produce goods in industry 2, and there may be some natural association between technology and markups. When $\lambda_n > \lambda_s$, it is almost trivial that $\frac{X_{n2}}{X_{n1}} > \frac{X_{s2}}{X_{s1}}$, but since the effect of the exchange rate is mainly a price one, the result that the South's export structure moving closer to the North's should remain similar, at least in the short run.

relevant measures: an index developed by Hausmann, Hwang, and Rodrik (2007) called *PRODY* that measures how heavily a good is exported by developed countries (see Section 2.3.1 for more details of this measurement) and a good's estimated elasticity of substitution by Broda and Weinstein (2006). Figure 2.1 shows a nonparametric relationship between the elasticity of substitution that we obtain from Broda and Weinstein (2006) and the export similarity index used in our empirical analysis. Clearly, there is a fairly strong negative correlation between these two.[29]

2.3 Estimation Strategy

2.3.1 Data and Variables

Our study draws on data from two sources. The first one is the China Customs data from 2000 (the earliest year of the data) to 2006 (the most recent year the authors have access to). The data set is at firm-product-destination-month level, covering a universe of all monthly import and export transactions by Chinese exporters and importers. Specifically, it includes product information (HS 8-digit-level classification), trade value, identity of Chinese importers and exporters, and import and export destinations.

The second data source is the International Financial Statistics (IFS) maintained by the International Monetary Fund (IMF), from which we obtain the monthly bilateral nominal exchange rates between China and the U.S. for the 2000-2006 period.

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[29] Both PRODY and elasticity of substitution are at HS 3-digit level. The PRODY at HS 3-digit level is the trade-weighted average of PRODY at HS 6-digit level. The HS 3-digit elasticity of substitution is estimated based on U.S. trade data, and downloaded from http://www.columbia.edu/~dew35/TradeElasticities/TradeElasticities.html, see Broda, Greenfield, and Weinstein (2006). Moreover, the fitted curve excludes the top 5% sigma, i.e. 7 sigmas with a value greater than 10.

To characterize China's export structure to the U.S., we first construct an index that differentiates each export product. Specifically, we use the measurement developed by Hausmann, Hwang, and Rodrik (2007), i.e.,

$$PRODY_i = \frac{1}{N_i}\sum_j \frac{X_{ij}}{X_j} GDPPC_j,$$

where X_{ij} is the export value of good i by country j; X_j is country $j's$ total export value; $GDPPC_j$ is the real per capita GDP of country j; and N_i is a normalization term used to have the coefficients summed up to 1. The intuition behind this measurement is that a good with a higher value of $PRODY_i$ is exported more often by developed countries.

In the empirical analysis, we use COMTRADE data to compute $PRODY_i$ for each HS-6 product in 2000 (the initial year of our data),[30] and then use the China Customs data to obtain a measure of overall export structure Y_{cm} (denoted as Export Similarity Index) for each city c in each month m during the period of 2000-2006, i.e.,

$$Y_{cm} = \sum_i PRODY_i \frac{X_{icm}}{X_{cm}},$$

where X_{icm} is the export value of good i to the U.S. by city c at month m; and X_{icm} is the total export value to the U.S. by city c at month m.

By fixing $PRODY_i$ in the initial year, we attribute the change in the city-level measurement Y_{cm} to the change in the allocation of exports across different product categories (i.e., changes in $\frac{X_{icm}}{X_{cm}}$). In other words, this

[30] Results remain robust when we measure $PRODY_i$ in other years before currency appreciation.

approach allows us to capture the change in the export structure, specifically, the similarity of export structure between China and developed countries.

To get a sense of $PRODY_i$, we list in Table 2.1 the five HS-6 product categories with the lowest values of $PRODY_i$ and the five HS-6 product categories with the highest values. Consistent with our intuition, goods with the lowest values of $PRODY_i$ are largely agricultural products, such as "Vegetable products nes", "Sisal and Agave (raw)", and "Cloves (whole fruit, cloves, and stems)". In the meantime, goods with the highest values of $PRODY_i$ are mostly metallic goods, such as "Cermets and articles thereof (waste or scrap)", "Sections H iron or non-alloy steel (nfw hot-roll/drawn/extruded > 80m)", "Sheet piling of iron or steel", and "Flat-rolled iron or non-alloy steel (coated with aluminium, width > 600mm)".

An alternative measurement of export structure in the literature is the one proposed by Schott (2008), based on Finger and Kreinin (1979)'s export similarity index (ESI). Specifically, it calculates the similarity between China's export structure and those of some developed countries (such as OECD countries), and the higher values mean more similarity. To calculate this measure, we need export data from other developed countries, which are available to us at the yearly frequency (i.e., via UN's COMTRADE data). However, our identification requires a measure at the monthly level. Nonetheless, we find that the yearly correlation between the export similarity indices developed by Hausmann, Hwang, and Rodrik (2007) and by Schott (2008) is 0.859, suggesting a robustness of using the former measure.

2.3.2 China’s Exchange-Rate Reform in July 2005

Timeline. Since the financial crackdown in 1994, China had adopted a decade-old fixed-exchange-rate regime, in which her currency (Renminbi) was pegged to the U.S. dollar at an exchange rate of 8.28. At 19:00 on July 21, 2005 (Beijing time), the People's Bank of China (PBOC, the central bank of China) suddenly announced a revaluation of Chinese currency against the U.S. dollar, which was set to be traded at an exchange rate of 8.11 immediately or about 2.1% appreciation. After that, Renminbi was allowed to trade flexibly with a reference basket of currencies with the target for Renminbi set by the PBOC every day. Figure 2.2 displays the monthly exchange rate between Chinese currency and the U.S. dollar during 2000-2006. It is clear that there was a sudden appreciation of Chinese currency against the U.S. dollar in July 2005, followed by steady and continuous appreciation. By the end of 2006, Renminbi had accumulated appreciation of about 5.5% against the U.S. dollar.

Exogeneity. The crucial part of our identification is to use the currency appreciation in China in mid-July 2005 as an exogenous shock; hence, it is important to establish the exogeneity of China's currency appreciation upfront. Note that the revaluation of Chinese currency in mid-July 2005 happened during a period of enormous international pressures on the Chinese government to appreciate her undervalued currency. However, the exact timing of the change has been widely considered as "unexpected". There is much anecdotal evidence as well as academic studies supporting this statement.

First, foreign pressures on Renminbi appreciation had existed for more than two years, and the Chinese government regarded the exchange-rate policy as a matter of China's sovereignty and rejected any political pressures on this issue. For example, on June 26, 2005 (about a month before the currency revaluation), China's Premier Wen Jiabao said at the Sixth Asia-Europe Finance Ministers Meeting in Tianjin that China would "independently determine the modality, timing, and content of reforms" and rejected foreign pressures for an immediate shift in the nation's currency regime.[31] One day later, Zhou Xiaochuan, the governor of the PBOC, said that it was too soon to drop the decade-old fixed-exchange-rate regime and that he had no plans to discuss the currency issue at the weekend meeting of the global central bankers in Basel, Switzerland.[32] On July 19, two days before the reform, the PBOC still insisted that it would continue to keep the exchange rate stable and at a reasonable and balanced level in the second half of the year.[33]

Second, as elaborated by Yuan (2012), opinions were divergent among Chinese policy makers regarding whether Chinese currency should be appreciated during that period. Specifically, the Ministry of Commerce opposed the currency appreciation (so as to maintain the competitiveness of China's export sector), while the other three central governmental agencies: the PBOC, the National Development and Reform Commission, and the Ministry of Finance, all proposed to appreciate Chinese currency.

Third, after the reform, both domestic and international medias responded to the revaluation with complete surprise. For example, CNN reported the

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[31] See "Chinese Premier Warns against Yuan Reform Haste" by the Wall Street Journal (http://online.wsj.com/news/articles/SB111975074805069620) Access date: October 9, 2012.

[32] See "China's Zhou Says `Time Is Not Ripe' to Drop Yuan Peg to Dollar" by Bloomberg (http://www.bloomberg.com/apps/news?pid=newsarchive&sid=aVAXsXEqKZcY&refer=home) Access date: October 9, 2012.

[33] See "China to Keep RMB Exchange Rate Basically Stable: Central Bank" by People's Daily (http://english.peopledaily.com.cn/200507/20/eng20050720_197148.html) Access date: October 9, 2012.

episode as "The surprise move by China...".[34] In the Financial Times' famous Lex Column on July 22, 2005 it was reported that "China likes to do things [in] its own way. After resisting pressure to revalue the Renminbi for so long, Beijing has moved sooner than even John Snow, the U.S. Treasury secretary, expected".[35] On July 22, 2005 the BBC Worldwide Monitoring said that "The People's Bank of China [PBOC] unexpectedly announced last night that the RMB [Renminbi] will appreciate by 2 per cent and will no longer be pegged to the U.S. dollar".[36]

Fourth, academic studies also imply that the change in the exchange-rate policy in July 2005 was unexpected. For example, Eichengreen and Tong (2011) study the impact of the Renminbi revaluation announcement on firm value in the 2005-2010 period. Using the change of stock prices before and after the announcement of the revaluation for 6,050 firms in 44 countries, they find that Renminbi appreciation significantly increases firm values for those exporting to China while significantly decreases firm values for those competing with Chinese firms in their home markets, suggesting the exogeneity of the policy change.

2.3.3 Estimation Framework

To identify the effect of currency appreciation on export structure, we exploit the sudden and unexpected currency revaluation by the Chinese government

[34] See "World Events Rattle Futures" by CNN (http://money.cnn.com/2005/07/21/markets/stockswatch/index.htm) Access date: October 9, 2012.

[35] See "Renminimal THE LEX COLUMN" by Financial Times (http://www.lexisnexis.com.libproxy1.nus.edu.sg/ap/academic/) Access date: October 9, 2012.

[36] See "Hong Kong Daily Says Exchange Rate Reform Advantageous Overall" by BBC Worldwide Monitoring (http://www.lexisnexis.com.libproxy1.nus.edu.sg/ap/academic/) Access date: October 9, 2012.

on July 21, 2005. Specifically, the unexpectedness in the currency revaluation makes the export structure before the revaluation a good counterfactual to the one after the revaluation. In other words, the exogenous currency appreciation in China offers us a regression discontinuity (RD) setting, which is arguably closest in the observational data analysis to the experimental design (e.g., Lee and Lemieux, 2010).

Hahn, Todd, and Van der Klaauw (2001) show that the RD estimator (β) can be identified as

$$\beta = \lim_{m \downarrow m_0} E[y_{cm} \mid m] - \lim_{m \uparrow m_0} E[y_{cm} \mid m],$$

where $y_{cm} = \ln Y_{cm}$; and $m_0 = July 2005$ is the cutoff month of the currency revaluation in China. Empirically, we focus on the data of the year 2005, use the local linear regression (as suggested by Hahn, Todd, and Van der Klaauw, 2001) with the triangle kernel function and the optimal bandwidth selected based on the procedure by Imbens and Kalyanaraman (2012), and obtain standard errors through the bootstrapping method.

However, there are two potential concerns about the above RD estimator. First, it may also capture the seasonal effect. For example, it could be that demand in the U.S. is different between July and August, causing the composition of Chinese exports to the U.S. to be different in these two months. In other words, $\hat{\beta}_{RD}$ becomes $\beta + \theta_{month}$, where θ_{month} is the monthly effect of exports. Second, the RD estimator essentially compares China's export structure to the U.S. in August 2005 with that in July 2005. Hence, one may be concerned whether the appreciation effect can be realized within such a short time window, especially given some pre-existing procurement contracts and the complexity of production.

To address these concerns, we include data of a year during which Chinese currency was fixed against the U.S. dollar, as a control group. Specifically, we choose the year 2003 as the month of Chinese New Year was the same for 2003 and 2005 (i.e., February), but different between 2004 and 2005 (i.e., January in 2004). Assuming the monthly effect is the same for these two years, we use a DD analysis to isolate the currency appreciation effect from the monthly effect, i.e.,

$$y_{cmt} = \alpha_t + \beta \cdot Aug_m \times Y2005_t + \psi_m + \varepsilon_{cmt}, \tag{2.4}$$

where $t \in \{2003, 2005\}$ represents year; α_t is the year fixed effect; $Aug_i = I[m \geq m_0]$ is an indicator of post-appreciation month;[37] $Y2005_t = I[t = 2005]$ is an indicator of the year 2005; and ψ_m captures the monthly effect. The standard errors are clustered at the month level, following Bertrand, Duflo, and Mullainathan (2004).

In addition to purging the monthly effect, the DD estimator, by comparing the five-month average export structure in the post-appreciation period with the seven-month average in the pre-appreciation period, reasonably captures the short-term effect of currency appreciation on export structure.

Note that Equation (2.4) uses an unbalanced city-level sample without inclusion of city fixed effects. Hence, we are estimating the overall effect of

[37] Empirically, we round m_0 to August as we only observe monthly trade data. Nonetheless, defining that the month of July has a value of 1/3 produces similar estimates (results available upon request).

currency appreciation on export structure. Later, we will experiment with the regression with the inclusion of city fixed effects (which captures the within-city and across-firms effect of currency appreciation) and the regression using firm-level data with the inclusion of firm fixed effects (which captures the within-firm and across-products effect of currency appreciation).

2.4 Empirical Findings

2.4.1 Main Results

Table 2.2 reports our estimates of the currency-appreciation effect on the structure of Chinese exports to the U.S. Column 1 shows the RD estimate using data of the year 2005. It is found that currency appreciation has a negligible effect on export structure: the effect is -0.2% and statistically insignificant.

However, such estimates may be biased due to some seasonal effects, e.g., U.S. demand difference between July and August. Column 2 presents the DD estimate by using data of the year 2003 as a control group. The DD estimate becomes positive and statistically significant, implying that currency appreciation makes the structure of Chinese exports to the U.S. more similar to those by developed countries.

Figure 2.3 presents graphically the results corresponding to Table 2.2: the dots and crosses represent the mean value of the similarity of Chinese export structure with developed countries' (in logarithm) for each month in 2003 and 2005, respectively; whereas the fitted curves are calculated using local linear regression with triangle kernel function. Clearly, the export similarity value dropped significantly from July 2003 to August 2003, suggesting a strong monthly effect. Differencing out such monthly effects, there was a sizable increase in the similarity value between July 2005 and August 2005, consistent with our estimates in Table 2.2. Meanwhile, export similarity values from January to June in 2003 and 2005 followed quite parallel trends, lending support to the argument that currency revaluation in July 2005 was largely exogenous and data in 2003 constructs a good comparison group for data in 2005.

Note that our estimates of the appreciation effect on export structure could be underestimated due to at least two reasons. Firstly, despite the fact that the exact timing of currency revaluation (i.e., July 2005) was completely unexpected, there had been some expectation that the Chinese government might revalue her currency since mid-2004. Such an expectation may make some producers change their behavior (like product upgrading decisions) earlier than the occurence of currency appreciation, causing an underestimation of our effect of interest. Secondly, our DD estimator captures largely a short-term effect of currency appreciation. In the long run, producers can upgrade their technologies, acquire advanced management practices, and recruit intelligent employees, all of which make our estimate underestimated.

2.4.2 Robustness Checks

In this subsection, we present a battery of robustness checks on our aforementioned estimation results in Table 2.3.

Alternative way of controlling for the monthly effect. While the inclusion of the year 2003 data helps us control for the monthly effect arising from the U.S. market situation, one may be concerned that the economic environment in China changed from July 2005 to August 2005, which spuriously generates the positive relationship between currency appreciation and change in the export structure.[38] As a check on such concerns, we look at the structure of Chinese exports to Nigeria, a country whose currency remained stable against Chinese currency in 2003 and 2005 especially between July and August (see Appendix Figure 2.1 for details). Column 1 of Table 2.3 reports a DD estimate of Equation (2.4) using Chinese export data to Nigeria. It is found that the estimated coefficient is highly insignificant and the magnitude is close to zero, indicating no significant changes in the Chinese market situation at the time of currency revaluation.

DDD estimation. In column 2 of Table 2.3, we combine Chinese export data to the U.S. in 2003 and 2005 with Chinese export data to Nigeria in 2003 and 2005, and conduct a difference-in-difference-in-differences (DDD) estimation, which enables us to control for the monthly effect arising from changes in both Chinese and foreign markets. Clearly, we find an estimate of 0.017, similar to that in column 2 of Table 2.2 (i.e., 0.018), suggesting the robustness of our previous findings.

Placebo test – pre-revaluation period. Given that Chinese currency was pegged to the U.S. dollar in 2002 and 2003, there was no break in the exchange rate between July and August in these years. Meanwhile, tariff reduction in China happened in the beginning instead of in the middle of the year; as a result, tariff reduction shall not contaminate our estimation. Hence, a DD estimation using data of the year 2002 and the year 2003 shall generate

[38] Note that tariff reduction in China happened in the beginning instead of in the middle of the year; as a result, tariff reduction shall not contaminate our estimates.

zero appreciation effect. Indeed, we find, in column 3 of Table 2.3, the DD estimator is highly insignificant and its magnitude is close to zero.

U.S. exports to China. As the appreciation of Chinese currency against the U.S. dollar means the depreciation of the latter against the former, we shall expect a reversed sign using the U.S.'s export structure to China as the outcome variable. Column 4 of Table 2.3 reports the DD estimate using Chinese imports from the U.S. in 2003 and 2005. Consistently, we find a negative and statistically significant estimated coefficient, implying that the appreciation of Chinese currency against the U.S. dollar makes U.S. exports to China more similar to those exported by developing countries.

Exclusion of processing trade. A unique feature of the Chinese trade system is that China allows some firms to import intermediate inputs free of tariffs but to export all their output, the so-called processing trade regime (e.g., Yu, 2014). One may be concerned that our results are driven by this special trade regime, hence compromising the external validity of our findings. To address such concern, we, in column 5 of Table 2.3, focus on the analysis of ordinary exports. Evidently, we still find a positive and statistically significant effect of appreciation on export similarity towards developed countries' export structures. Meanwhile, despite a slight drop, the estimated magnitude (0.014) is statistically indifferent from the estimate in our benchmark model (i.e., 0.018 in column 2 of Table 2.2).

2.4.3 Decomposition of the Effect of Currency Appreciation

In this section, we use our data to decompose the resource reallocation at different margins (i.e., across cities, across firms within a city, and across products within a firm). Our previous analyses use an unbalanced city/firm sample; hence, the DD estimate in column 2 of Table 2.2 is the overall effect of currency appreciation, including resource reallocation across cities, within city and across firms, and within firm and across products. To decompose the currency appreciation effect on export structure into these three different margins, we conduct two more regressions in Table 2.4. Specifically, in column 1, we include city dummies and in column 2, we use a sample of surviving multi-product exporters (i.e., those that existed before and after currency revaluation) with an inclusion of firm dummies.[39]

Both coefficients are found to be positive and statistically significant, consistent with our previous findings. Meanwhile, as the analysis with the inclusion of city dummies essentially calculates the effect of appreciation on the within-city change in export structure, the comparison of the coefficient with the one without city dummies (i.e., the one in column 2 of Table 2.2) can give us the degree of across-cities resource reallocation effect of currency appreciation. Similarly, the comparison of coefficients between column 1 and column 2 can allow us to gauge the magnitude of within-city, across-firms resource reallocation effect of currency appreciation. Finally, the coefficient in column 2 produces the within-firm, across-products resource reallocation effect of currency appreciation.

It is found that the majority of the currency appreciation effect on export structure comes from the resource reallocation within city and across firms, i.e., accounting for (0.015-0.002)/0.018=72.22%. Meanwhile, the across-cities

[39] Note that we put the change in export revenues for single-product firms in the category of the resource reallocation within city and across firms.

resource reallocation accounts for around (0.018-0.015)/0.018=16.67%, while the within-firm, across-products resource reallocation accounts for 11.11%.

2.5 Conclusion

This paper investigates whether and how a country's export structure responds to its exchange-rate movement. Using China's sudden and unexpected revaluation of its currency against the U.S. dollar, we identify the effect of exchange rates on export structure through the combined regression discontinuity and difference-in-differences framework. We find that after its currency appreciation, China's exports to the U.S. became more similar to those by developed countries. Meanwhile, we find that the majority of the currency-appreciation effect on export structure comes from the resource reallocation within city and across firms.

The major implication of our empirical findings is that the depreciation strategy used by developed countries may reduce the import competition from developing countries, but not by much. Despite of the fact that our empirical results are necessarily short-term by the identification strategy, changes in export structure may have important long-run implications, especially with the resource reallocation and learning-by-doing effects.

Chapter Three

When Trade Discourages Political Favoritism: Evidence from China

3.1 Introduction

Since Adam Smith and David Ricardo, economists have espoused the benefits of free trade. A brief aberration occurred in the late 1970s and early 1980s, when the New Trade Theory shows that in a world with market imperfections, trade barriers could in fact increase national welfare (e.g., Brander and Spencer 1982; Krugman1982; Dixit 1984). However, as new research uncovers heretofore overlooked sources of gains from trade in recent years, there is a growing realization that the welfare gains from trade might have been underestimated after all (e.g., Melitz 2003; Lileeva and Trefler 2010; Melitz and Redding 2014).[40]

One previously overlooked source of trade gains is the increase in overall productivity when trade liberalization induces a reallocation of resources from less productive to more productive firms (Melitz 2003; Melitz and Ottaviano 2008). This channel through which trade enriches a nation appears to be particularly relevant to developing economies, where firms differ in productivities more than their counterparts in the developed world and there exists considerable room to improve allocative efficiency (Hsieh and Klenow 2009; Pages 2010). One factor that contributes to the greater dispersion of productivity in developing countries is government protection of inefficient firms. For instance, it is well-documented that state-owned enterprises (SOEs)

[40] See Melitz and Trefler (2012) for a review.

in China enjoy favorable access to resources such as cheap credit and land, [41] even though they are less efficient than non-SOEs.[42] If the government displays favoritism toward some firms over others before trade liberalization, the cost of providing such support may increase when trade liberalization induces more market competition. Any subsequent withdrawal of political favoritism would improve resource allocation and generate productivity gains. In this paper, we empirically investigate if there exist such gains from trade (generated through reducing political economy inefficiencies) by studying China's accession to the World Trade Organization. We examine if the event reduces the market share of the inefficient but politically favored SOEs.

Our analysis is based primarily upon the 1998--2005 Annual Survey of Industrial Firms (ASIF), the most comprehensive firm-level data in China. Replicating the strategy in Topalova (2007), we use China's accession to the WTO in December 2001 to conduct a difference-in-differences (DD) analysis on Chinese cities: our identification strategy exploits variations in city-level industrial composition, which generated differential trade shocks across cities after tariffs were lowered. This allows us to compare the SOE output and employment shares in cities that experienced larger degrees of trade liberalization with those that experienced smaller degrees of trade liberalization before and after China's WTO accession.

We find that trade liberalization significantly reduced both the output and employment shares of SOEs. In our preferred specification, trade liberalization

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[41] According to Liu and Zhou (2011), large and medium-sized non-SOEs face an average interest rate that is 6 percentage points higher than SOEs of corresponding size, whereas the average interest rate of small non-SOEs is 9 percentage points more than that of small SOEs.

[42] According to the Chinese Statistical Yearbook, the ROA of industrial SOEs was 3.0% in 2002, while the ROAs of foreign invested firms and domestic non-SOEs were 6.0% and 5.6% respectively. See also World Bank and Development Research Center of the State Council, P.R.C. (2012, ch. 3).

induced a 2.48 percentage points decline in SOE output share between 2001 and 2005, or 16.4% of the actual decline observed during this period.[43] Based on the efficiency gain estimates calculated by (Hsieh and Klenow (2009), we infer that trade liberalization through the WTO accession reduced the standard deviation of manufacturing firm productivity in China by 1.41%. While this reduction may appear modest, it is an additional welfare gain on top and above the traditional gains of trade arising from country differences and comparative advantage.

What drove the post-WTO accession contraction of SOE output and employment shares? Further investigation indicates that it was mainly driven by an increase in import competition instead of improved access to export markets or cheaper imported intermediate inputs. In addition, the contraction took place across a variety of industries and was not confined to the industries initially dominated by SOEs. Finally, like Brandt, Van Biesebroeck, Wang, and Zhang (2012), we find that the contraction occurred at the extensive margin (i.e., due to exit) instead of the intensive margin (i.e., due to surviving SOEs losing output share).

Interestingly, we find that SOEs affiliated to county and township governments were more likely to exit after China's WTO accession, while SOEs affiliated with higher levels of government (central, provincial, and city) were largely unaffected. In other words, the SOEs that exited the market after December 2001 were not the least productive ones—as existing theoretical models such as Melitz and Ottaviano (2008) would have predicted—but those

[43] The city-level average output share of Chinese SOEs fell from 29.55% in 2001 to 14.42% in 2005. Besides trade liberalization, other factors that contributed to this decline include SOE reform and the relaxation of FDI regulations. See Section 3.3 for a detailed discussion.

with the weakest political backing. Since the fiscal health of higher level Chinese governments were far superior to that of the counties and townships,[44] this finding provides further evidence that the increased cost of supporting inefficient firms after China entered the WTO contributed to the observed decline in SOE output and employment shares. It also suggests that while trade liberalization induces productivity gains by making government support of inefficient firms more costly, such gains are made at the margin and some inefficiency is likely to persist as long as the government has the financial ability and political will to continue providing support.

Our work is related to several strands of the literature. First, there is a recent revival of interest in the sources and magnitude of gains from trade. In a much cited paper, Arkolakis, Costinot, and Rodriguez-Clare (2012) show that for a variety of trade models which satisfy the constant elasticity of substitution (CES) restriction, the gains from trade can be pinned down by two parameters: the share of expenditure on domestic goods (λ) and the elasticity of imports with respect to variable trade costs (ε). An implication of this result is that the gains from trade may be modest. Indeed, Eaton and Kortum (2002) estimate that the US would suffer a welfare loss of only 0.8% if it moves to autarky in manufactures. More recent studies have focused on examining the pro-competitive effects of trade which are overlooked under the CES restriction (e.g., de Blas and Russ 2010; Arkolakis, Costinot, Donaldson, and Rodriguez-Clare 2012; Edmond, Midrigan, and Xu 2012; Holmes, Hsu, and Lee 2013) and our paper is an effort in this direction.

[44] Local governments in China shoulder 80% of all public expenditure responsibilities but receive only 40% of the tax revenues (World Bank and Development Research Center of the State Council, P.R.C. 2012, Figure 0.8).

Our paper also makes contact with studies that investigate the mitigating effect of trade on export distortions (Khandelwal, Schott, and Wei 2013), tax distortions (e.g., Konan and Maskus 2000), and labor market distortions (e.g. Krishna, Yavas, and Mukhopadhyay 2005 and Krishna and Yavas 2005). In particular, Khandelwal, Schott, and Wei (2013) find that upon the expiration of the Multifiber Arrangement (MFA) in 2005, new entrants in China—most of them non-SOEs—drove up the volume of Chinese textile and clothing exports while driving down their prices. According to their structural estimation, an improved allocation of export quotas accounts for 71% of China's overall gains from the expiration of MFA.

There is a growing literature looking at China's accession to the WTO and these studies overwhelmingly indicate that the benefits of WTO membership are positive for China. Chen, Ma, and Xu (2014) propose a generalized trade restrictiveness index and use it to confirm the WTO's effectiveness in removing tariff barriers in China, while Yu (2014) detects a positive impact of WTO-associated tariff reduction on the productivity of Chinese firms. Brandt, Van Biesebroeck, and Zhang (2012) document that between 1998 and 2007, the productivity of incumbent firms grew at a weighted average of 2.9%—8.0% annually. Exploring sectoral variations in tariff reduction upon the WTO accession, Brandt, Van Biesebroeck, Wang, and Zhang (2012) show that trade liberalization reduces firm prices and markups. Fan, Li, and Yeaple (2013) find that WTO accession led to an improvement in the quality of Chinese exports. According to Han, Liu, and Zhang (2012), accession to the WTO significantly increased its wage inequality but much of this effect was due to an increase in returns to education. Using cross-sectional and panel data, Lan and Li (2013) show that trade weakens nationalism in China.

Finally, our study contributes to the literature on SOEs in China. Song, Storesletten, and Zilibotti (2011) show that the presence of inefficient but politically favored SOEs helps create the puzzling coexistence of high returns on capital and a growing foreign surplus in China. Du, Lu, Tao, and Yu (2014) argue that SOEs are costly to the Chinese economy in at least two dimensions: not only do they have lower production efficiency, they also possess higher market power than non-SOEs. Li, Liu, and Wang (2012) find that the improved performance of SOEs in recent years is not driven by a genuine improvement in efficiency but by the consolidation of a vertical industry structure whereby the SOEs monopolize key upstream industries while the non-SOEs compete in downstream industries. Likewise, Tang, Wang, and Wang (2014) show that SOEs register significantly higher ratios of domestic value added in exports than foreign invested firms and large domestic non-SOEs and they attribute this finding to the vertical industry structure in China.

The rest of the paper is organized as follows. Section 3.2 discusses our estimation strategy in detail. In Section 3.3, we present our empirical findings and conduct robustness checks. Section 3.4 concludes.

3.2 Estimation Strategy

3.2.1 China’s WTO Accession

In July 1986, China notified the GATT (the predecessor of the WTO) that it would like to resume its status as a GATT contracting party. Between 1987 and 1992, as China was debating the direction of its economic reform domestically, the return to GATT was suspended. The momentum resumed after Deng Xiaoping's southern tour speech in 1992, and in July 1995, China officially filed its application to join the WTO.

The pivotal part of China's WTO accession process involved bilateral negotiations between China and the existing members of the WTO. The first country that signed a bilateral WTO accession agreement with China was New Zealand (in August 1997). The negotiation between China and the U.S. was the toughest. It took the two countries four years and twenty-five rounds of negotiation before an agreement was reached in November 1999. Subsequently, China reached agreements with 19 countries within half a year, including Canada in November 1999 and the European Union in May 2000. In September 2001, China concluded the agreement with Mexico, which marked the completion of negotiations with all WTO member countries. Finally, the WTO's Ministerial Conference approved by consensus the text of the agreement for China's entry into the WTO on November 10, 2001.

To illustrate its commitment to join the WTO, China cut tariffs substantially between 1992 and 1997. In 1992, China's (unweighted) average tariff rate was as high as 42.9%. Shortly after the GATT Uruguay round negotiations, China lowered tariffs from an average rate of 35% in 1994 to 17% in 1997. Tariff rates remained stable after 1997 until China officially joined the WTO on December 11, 2001. From 2002 onward, China took steps to fulfil her tariff reduction responsibility as a WTO member country. According to the accession agreement, China would fulfil its promised tariff cuts by 2004 (with a few exceptions to be completed by 2010) and the average tariff rates for agricultural and manufacturing products would be reduced to 15% and 8.9% respectively.

Figure 3.1 plots China's (unweighted) average tariffs for the period 1996—2007. It shows that the tariff rates experienced a substantial drop in 1996. This was followed by a relatively stable period between 1997 and 2001 and another round of gradual cuts in 2002, before a steady state was reached in 2005. The unweighted average tariff rate dropped from 15.3% in 2001 to 12.3% in 2004 while the weighted average tariff rate fell from 9.1% to 6.4%.

Furthermore, the dispersion of tariffs was significantly reduced after China's WTO accession. As shown in Figure 3.1, the ratio of tariffs at the 25th percentile over those at the 75th percentile experienced a sharp drop in 2002 and stabilized only after 2005. In Figure 3.2, we plot the relationship between tariff rates in 2001 and tariff rate changes between 2001 and 2005 across three-digit industries (the unit which we use to construct the city-level exposure to trade liberalization; see Section 3.2.3 for details).[45] We observe a strong, positive correlation, implying that industries with higher tariffs before China's WTO accession experienced more tariff reduction after that. This is perhaps unsurprising since China was free to set different tariffs for different industries before 2001 and this freedom was lost when it became a WTO member and had to reduce tariff rates to the WTO-determined levels which are relatively uniform across products.

[45] We find similar patterns at the HS-6 product level (results available upon request).

3.2.2 Data

The main dataset used in this study comes from the Annual Survey of Industrial Firms (ASIF), conducted by the National Bureau of Statistics (NBS) of China for the period of 1998—2005. It is the most comprehensive firm-level dataset in China.[46] The data contain all state-owned enterprises (SOEs) and non-SOEs with annual sales of five million RMB (around US$600,000) or more. The number of firms varies from over 140,000 in the late 1990s to over 243,000 in 2005, spanning all 31 provinces or province-level municipalities (covering 344 cities and 2,829 counties) and all manufacturing industries (29 two-digit, 164 three-digit and 464 four-digit industries).[47] The dataset provides detailed firm-level information, including firm name, industry affiliation, location, and all operation and performance items reported in accounting statements such as age, employment, capital, intermediate inputs, and ownership.

Our outcome variable concerns the comparative performance of SOEs and non-SOEs, which requires us to first identify SOEs in our sample. In the benchmark analysis, we follow the official definition of SOEs in the data.

[46] This dataset is noted for its representativeness because the firms sampled contribute the bulk of China's industrial value-added output. The dataset is used to calculate key national income accounting metrics (e.g., GDP) and other statistics published in China's official statistical yearbooks. This dataset is found to be reasonably accurate and reliable due to strict double-checking procedures in the data collection process (Cai and Liu 2009). Thus, it has been widely used by economic researchers in recent years (Bai, Lu, and Tao 2009; Cai and Liu 2009; Lu, Lu, and Tao 2010; Brandt, Van Biesebroeck, and Zhang 2012).

[47] During the period sampled, there were some adjustments in China's administrative boundaries. In some cases, new counties were established. In others, existing counties were merged to form larger counties or cities. To maintain consistency in our coding of cities and counties, we use the 1999 National Standard (promulgated at the end of 1998 and known as the GB/T 2260-1999) as the benchmark codes and convert the regional codes of all firms to these benchmark codes. Separately, in 2003 a new classification system for industry codes (GB/T 4754-2002) replaced the old classification system (GB/T 4754-1994) in use from 1995 to 2002. To maintain consistency in our coding of industries for the entire period sampled (1998—2005), we convert the industry codes in the 1998—2002 data to the new classification system.

Specifically, according to the NBS categorization, SOEs correspond to three specific registered ownership types in our data: (a) code 110, state-owned enterprises; (b) code 141, state-associated enterprises; and (c) code 151, enterprise solely funded by the state. As a robustness check (Appendix Table 3.1, column 1), we use an alternative definition of SOEs proposed by Hsieh and Song (2013), who classify a firm as a SOE if it satisfies one of the two following conditions: (a) the registered capital held directly by the state exceeds 50%; or (b) the ASIF data identifies the state as the controlling shareholder of the firm.

The dataset of Chinese tariffs is downloaded from the WTO website. Specifically, we use the *Tariff Download Facility* to obtain the standardized tariff statistics. For each product defined at the HS-6 digit level, the tariff data provide detailed information including the number of tariff lines, the average, minimum, and maximum ad valorem tariff duties. The tariff data is available for 1996, 1997 and 2001 (latest). As the WTO website does not provide tariff information for 1998—2000, we use the World Integrated Trade Solution website maintained by the World Bank to fill the void. Meanwhile, as different HS codes are used before and after 2002, we match the 1996 HS codes (used in the 1997—2001 tariff schedules) to the 2002 HS codes (used in the 2001—2006 tariff schedules) using the standard HS concordance table. Furthermore, as the ASIF data is classified at the industry-level, we need to aggregate tariffs from the HS-product level to the industry-level. To this end, we first match the HS classification to the Chinese Industrial Classification (CIC) using the concordance table from the National Bureau of Statistics of

China.[48] Subsequently, we calculate the simple average tariff for each industry and each year.

Finally, in some parts of our analysis, we include several city-level characteristics based on the Chinese City Statistical Yearbook (multiple years). These variables include GDP, population, government consumption, vegetable consumption, dairy consumption, number of telephones, and number of colleges.

3.2.3 Estimation Specification

To examine the differential impacts of trade liberalization on SOEs and non-SOEs, we follow the locality-event difference-in-differences (DD) approach devised by Topalova (2007).[49] Specifically, because the geographic location of industrial activities varied across Chinese cities before China's WTO accession, the sudden tariff reduction upon accession generates differential impacts on the cities. This allows us to identify the effect of trade liberalization on SOEs.

We conduct the analysis at the city-level instead of the industry-level for two reasons. First, generally speaking, SOEs in China are affiliated with territorial administrative units (i.e. center, province, city, county, or township) instead of functional units (i.e. by ministry or industry). Second, a city-level analysis allows us to capture the general equilibrium effect of trade liberalization on SOEs' activities, for example, trade liberalization may affect

[48] We thank Yifan Zhang for sharing this concordance table.

[49] For studies applying this identification strategy, see Hasan, Mitra, and Ural (2007); Edmonds, Pavcnik, and Topalova (2010); McLaren and Hakobyan (2010); Topalova (2010); McCaig (2011); Hasan, Mitra, Ranjan, and Ahsan (2012); Autor, Dorn, and Hanson (2013). See Kovak (2013) for a micro-foundation of this identification strategy.

prices of local tradable and non-tradable goods as well as local wages and employment rates. In the latter part of the paper, we will look at whether the effect of trade liberalization comes from within-industry reallocation (i.e., a decline of SOEs' activities within an industry) or cross-industry reallocation (i.e., a contraction of SOE-dominated industries relative to industries with smaller SOE presence).

The specification of our DD estimation is

$$y_{ct} = \alpha_c + \beta Tariff_{ct} + X_{ct}^{'}\gamma + \left(Z_{c2001} \cdot Post02_t\right)^{'} \cdot \theta + \lambda_{pt} + \varepsilon_{ct}, \qquad (3.1)$$

where c and t represent city and year, respectively, and ε_{ct} is the error term. To deal with potential heteroskedasticity and serial autocorrelation, we cluster the standard errors at the city level (as recommended in Bertrand, Duflo, and Mullainathan 2004).[50]

α_c is the city fixed effect, controlling for all time-invariant differences across cities; λ_{pt} is the province-year fixed effect, controlling for all annual shocks common to cities (such as business cycles, macro policies, etc.) and all provincial heterogeneity (including all time-invariant and time-varying characteristics); and X_{ct} is a vector of time-varying city characteristics (including GDP per capita and the share of government consumption) that are potentially correlated with both our outcome variable and the regressor of interest and are thus included to isolate the trade liberalization effect.

Our outcome variable, y_{ct}, measures the share of SOEs in city c at year t. In the benchmark analysis, we focus on the output share of SOEs over all

[50] We also use another approach devised by Bertrand, Duflo, and Mullainathan (2004), where we collapse the panel structure into two periods (i.e., one before and the other after China's WTO accession), and use the White-robust standard errors. We obtain similar results (see Appendix Table 3.1, column 2).

firms instead of the employment share due to labor hoarding issues in SOEs. Nonetheless, using employment share as the outcome variable generates similar results (see Appendix Table 3.1, column 3). There are three potential concerns with the SOE share measurement, and it is worthwhile to discuss them upfront. First, the ASIF data is truncated as small non-SOEs (i.e., those with annual sales below 5 million RMB) are not sampled. Hence, if trade liberalization results in both small SOEs and non-SOEs exiting the market, we would mistakenly detect a stronger trade effect on SOEs than on non-SOEs due to the truncated nature of the data. To address this measurement concern, as a robustness check we exclude SOEs with annual sales below 5 million RMB in our outcome variable (see Appendix Table 3.1, column 4). Second, when calculating the output share of SOEs, we average over all firms including SOEs, domestic non-SOEs, and foreign-invested firms. It is possible that an observed fall in the output share of SOEs could be driven by a surge in the sales of foreign firms that is not at the expense of the SOEs. To address this potential concern, we conduct a robustness check that excludes the sales from foreign firms when calculating the output share of SOEs (see Appendix Table 3.1, column 5). Third, when calculating the output share, we include output sold in the domestic market as well as output sold overseas. To the extent that tariff reduction affects mostly domestic competition, one may be concerned that the inclusion of foreign sales may potentially bias our results. As a robustness check, we exclude foreign sales in our outcome variable (see Appendix Table 3.1, column 6).

The regressor of interest, $Tariff_{ct}$, captures the city-level exposure to trade liberalization. Specifically, it is measured as

$$Tariff_{ct} = \frac{\sum_i Output_{ic2001} \times Tariff_{i2001} \times Post02_t}{\sum_i Output_{ic2001}}, \tag{3.2}$$

where i represents the manufacturing industry; $Output_{ic2001}$ is the total output of industry i in city c in 2001; $Tariff_{i2001}$ is the import tariff rate of industry i in 2001; and $Post02_t$ indicates the period when China becomes a WTO member, taking the value of 1 for year 2002 or after, and 0 otherwise. As a robustness check, we replace $Tariff_{i2001}$ with the average tariff rate of industry i in 1997—2001 (i.e., $Tariff_{c1997-2001}$) and find similar results (see Appendix Table 3.1, column 7).

There are two issues about our tariff measurement worth pointing out upfront. First, we use the interaction of tariffs in 2001 ($Tariff_{i2001}$) and the post-WTO accession indicator ($Post02_t$) instead of yearly tariffs ($Tariff_{it}$), as used in Topalova (2007), for three reasons: (a) since China released its schedule of tariff reduction upon WTO accession in 2002 and firms could exploit this information, $Tariff_{it}$ is less exogenous than $Tariff_{i2001} \times Post02_t$; (b) since industries with higher tariffs in 2001 experienced more tariff reduction upon WTO accession (as illustrated in Figure 3.2), $Tariff_{i2001} \times Post02_t$ allows us to capture the effects of differential tariff reductions; and (c) as elaborated in Liu and Trefler (2011), using the interaction between $Tariff_{i2001} \times Post02_t$ allows us to capture not only the realized effects of trade liberalization across the period of study, but also the effects of unrealized (but anticipated) tariff cuts scheduled to be phased out after 2005. This helps to ensure that we are not overestimating the effects of trade liberalization. Second, following Topalova (2007), we ignore the nontraded sectors when calculating city-level tariffs. Kovak (2013) gives a justification of using such nonscaled tariff measurement; it avoids the estimation bias coming from the assumption that nontraded prices are unaffected by trade liberalization.

Z_{c2001} represents determinants of the geographic distribution of industrial activities across Chinese cities in 2001. It is interacted with the post-WTO accession indicator ($Post02_t$) to control for the differential effects of these potential preexisting differences between treatment and control groups on the SOE share (for a similar practice, see Gentzkow 2006). Ellison and Glaeser (1999) have characterized a list of determinants of geographic concentration, and we follow their approach closely by locating all determinants that they identified and are available in the Chinese City Statistical Yearbook. Specifically, we use the number of dairy consumption per capita, the number of vegetable consumption per capita, the number of telephones, and the number of universities in 2001, and measured in the logarithm form (except for the number of universities, as many cities have no universities). Furthermore, we categorize cities into northern cities (vs. southern cities), coastal cities (vs. inland cities), and mountain cities (vs. plain cities) to control for any differential geographic impacts. In addition, we also include city-specific linear time trend (i.e., $\alpha_c \cdot t$) as a regressor to control for the underlying differences between our treatment and control groups in a restricted way, that is, under the assumption that these preexisting city-level differences affect our outcome variable linearly with time.

We estimate Equation (3.1) using first differences, which removes the city fixed effect and reduces the degree of serial autocorrelation, i.e., we estimate

$$\Delta y_{ct} = \beta \Delta Tariff_{ct} + \Delta X'_{ct} \gamma + \Delta (Z_{c2001} \cdot Post02_t)' \cdot \theta + \lambda_{pt} + \Delta \varepsilon_{ct}, \qquad (3.3)$$

where Δ is the first-differenced operator, e.g., $\Delta y_{ct} = y_{ct} - y_{ct-1}$.

In addition, to provide further support on the validity of our DD specification, we conduct several robustness checks: the addition of controls for the geographic distribution of industrial activities; a flexible estimation to examine whether the treatment and control groups were comparable until the time of WTO accession; the addition of controls for other ongoing policy reforms in the early 2000s (i.e., the SOE reform and the relaxation of FDI regulations); a placebo test using only the pre-WTO accession period data as in Topalova (2010); and a placebo test using the sample of pure exporters (those that exported 100% of their output and were therefore unaffected by tariff reduction). For details, see Section 3.3.2.

3.3 Empirical Findings

The regression results of the DD specification (3.3) are reported in column 1 of Table 3.1. We find that the coefficient estimate of our regressor of interest (T$Tariff_{ct}$) is negative and statistically significant, suggesting that cities with higher effective tariffs in 2001 experienced a larger decline of SOE output share after 2002 than those with lower effective tariffs. Given that cities with higher tariffs in 2001 experienced more tariff reduction after 2002, these results imply that trade liberalization reduces SOEs' activities.

One potential concern of our DD estimation is that the geographic distribution of industrial activities across Chinese cities in 2001 may not be random. If this is true, cities with different industrial structures (and hence facing different effective tariffs) might be systematically different before WTO accession, and such preexisting differences may generate the spurious negative relationship between trade liberalization and SOE output share in the post-WTO accession period. To address this concern, in column 2 of Table 3.1, we include an interaction term comprising the post-WTO accession period indicator and the determinants of geographic concentration identified by Ellison and Glaeser (1999) where data are available. The coefficient estimate of our regressor of interest remains negative and statistically significant. In fact, the magnitude and statistical significance both increase, suggesting that any bias caused by these preexisting differences has led to an underestimation of the effect of trade liberalization.

3.3.1 Magnitude and Gains Calculation

To gauge the economic magnitude of our estimates, we conduct the following exercise: The mean values of city-level tariffs in 2001 and 2005 are 17.83% and 9.66% respectively. Hence, the predicted change in SOE output share from 2001 to 2005 is -(17.83%-9.66%)$*$0.304=-2.48 %, where 0.304 is the coefficient estimate of interest (Table 3.1 column 2). Meanwhile, during the 2001—2005 period, the actual mean value of SOE output share fell from 29.55% to 14.42%. Hence, trade liberalization can account for 2.48%/(29.55%-14.42%)=16.42% of the total change in SOE output share.

Next, we employ the methodology developed by Hsieh and Klenow (2009) using the dispersion of revenue-based TFP as a proxy to capture market distortion. Applying the method to the ASIF data (the same data as ours), they find that the standard deviation of revenue-based TFP in China in 2001 is 0.68. We replicate their estimation specification at the city level (instead of the industry level) and find that every 1% decrease in the SOE output share is associated with a 0.57% decrease in the standard deviation of revenue-based TFP.

Since the reduction of tariffs from 2001 to 2005 led to a reduction of the SOE output share by 2.48%, we can infer that the standard deviation of revenue-based TFP is reduced by 2.48%∗0.57=1.41%. In other words, the WTO-induced trade liberalization contributes to a 1.41% reduction in allocative inefficiency.

3.3.2 Robustness Checks

In this subsection, we report results of a battery of robustness checks on our aforementioned DD estimation.

Flexible estimation of treatment effect parameters. To check the comparability between our treatment and control groups, we conduct a flexible estimation specification; that is, we replace the post-WTO accession period indicator ($Post02_t$) in the construction of city-level tariff variable with year dummies (λ_t), so that the regressor of interest becomes $FlexTariff_{ct} = \frac{\sum_i Output_{ic2001} * Tariff_{i2001}}{\sum_i Output_{ic2001}} \times \lambda_t$. Figure 3.3 plots the estimated coefficients as well as the 95% confidence intervals from this exercise. It shows that in the pre-WTO accession period (1998—2001), our treatment and control groups have comparable time trends, as the coefficients stay relatively constant over time. This alleviates any concern that our treatment and control groups are ex ante incomparable and lends support to our DD identifying assumption. Meanwhile, there is a visible divergence between the two groups in their SOE output share trends after 2002, when China took steps to reduce its tariffs to honor its obligations to the WTO. The consistency in timing suggests that trade liberalization reduces the output share of SOEs.

City-specific linear time trend. Although we have controlled for the post-WTO accession time trend of SOE output share generated by the pre-WTO accession determinants of geographic concentration, one may still be concerned over some unobserved city characteristics omitted from the equation, which could compromise the comparability between our treatment and control groups. To check whether our estimates are biased due to these unobserved city factors, we include the city-specific linear time trend, i.e., $\alpha_c \cdot t$ (in the first-differenced equation, $\alpha_c \cdot t$ collapses to α_c, i.e., the city fixed effects). This additional control would allow us to control for all unobserved city characteristics if these characteristics affect our outcome variable in the specification of a linear time trend. Regression results are reported in column 1 of Table 3.2. Our regressor of interest remains negative and statistically significant, implying that our estimates are not driven by unobserved underlying city characteristics.

Control for other policy reforms. China adopted several reforms that overlapped with its WTO accession. First, in the early 2000s, it lifted some restrictions on FDI by increasing the number of industries that FDI is permitted and relaxing the constraints on local ownership requirements. To control for this reform, we add the number of foreign-invested firms (in logarithm) as a regressor in column 2 of Table 3.2. Our trade liberalization effect remains robust.

Second, China initiated a round of SOE reform in the late 1990s, which was still ongoing in the early 2000s. To control for this, we conduct three exercises. First, we include the percentage of SOEs that were being privatized as an additional control in column 3 of Table 3.2 and find that our findings remain robust. Second, we focus on a subsample of firms that did not experience a change in ownership status (i.e., they were either SOEs or non-SOEs throughout the period we study). We obtain similar results (reported in column 4 of Table 3.2). Finally, we use the degree of privatization as the outcome variable in column 5 of Table 3.2 and find that it is barely affected by trade liberalization.

Placebo test I: Pre-WTO accession period. As the first placebo test, we follow Topalova (2010) in looking at the effect of tariff changes on the SOE output share in the pre-WTO accession period (i.e., 1998—2001). Since the tariff schedule did not change significantly during this period, we expect a muted effect; otherwise, it may indicate the existence of some underlying confounding factors. As shown in column 6 of Table 3.2, we indeed find that tariff changes have no significant effect on the SOE output share in the pre-WTO accession period.

Placebo test II: Subsample of pure exporters. In our data, there are firms exporting 100% of their outputs. Since these pure exporters are not affected by domestic competition, tariff reduction upon WTO accession shall have a limited effect on them. We report the regression results in column 7 of Table 3.2. As expected, we find that the coefficient of $Tariff_{ct}$ is highly insignificant.

3.4 Mechanism

In the previous section, we establish that trade liberalization (induced by accession to the WTO) substantially reduced the output and employment shares of SOEs in China. To shed light on the underlying mechanisms, we first examine whether the effect of trade liberalization comes from the import competition channel, the exporting market access channel, or the imported intermediate inputs channel. Next, we investigate whether the decline in SOE activities is due to within-industry reallocation (i.e., decline of SOE output share within each industry) or cross-industry reallocation (i.e., the shrinkage of SOE-dominated industries). We then decompose the trade effect into the extensive (i.e., entry and exit) and intensive (i.e., output changes of surviving firms) margins. Finally, we investigate whether and how different SOEs respond to trade liberalization differently.

3.4.1 Import Competition, Export Market Access, Imported Inputs

The WTO pursues a multilateral and multidimensional agenda. As such, China's accession to the WTO involved not only China reducing its tariffs on manufactured imports, but also existing WTO member countries lowering their tariffs on Chinese exports. Furthermore, Chinese firms may also gain access to cheaper international inputs (see, e.g., Goldberg, Khandelwal, Pavcnik, and Topalova 2010). If China's export tariffs and input tariffs are perfectly correlated with its import tariffs (our regressor of interest), our aforementioned results should be interpreted as a general WTO effect instead of an import competition effect.

To differentiate these three channels (namely, increased import competition, better access to export markets, and cheaper imported intermediate inputs), we add as regressors China's export tariffs and input tariffs. Specifically, we measure the city-level exposure to export market by

$$Export\ Tariff_{ct}^{\mathrm{export}} = \frac{\sum_i Output_{ic2001} \times Tariff_{i2001}^{external} \times Post02_t}{\sum_i Output_{ic2001}}, \tag{3.4}$$

where $Tariff_{i2001}^{external} = \sum_f Tariff_{i2001} \times \frac{export_{fi2001}}{export_{i2001}}$; $Tariff_{fi2001}$ is foreign country f 's tariffs on Chinese imports of industry i in 2001; $export_{fi2001}$ is Chinese total exports of industry i to foreign country f in 2001; and $export_{i2001}$ is Chinese total exports of industry i in 2001. Meanwhile, we measure the city-level exposure to imported intermediate inputs by

$$Input\ Tariff_{ct}^{input} = \frac{\sum_i Output_{ic2001} \times Tariff_{i2001}^{input} \times Post02_t}{\sum_i Output_{ic2001}}, \tag{3.5}$$

where $Tariff_{i2001}^{input} = \sum_k Tariff_{k2001} \times \omega_{ki}$ and ω_{ki} is the share of inputs from industry k used by industry i, based on the 1997 Chinese input-output table.

The regression results are reported in columns 1–3 of Table 3.3. We find that neither export tariffs nor input tariffs are statistically significant. Furthermore, their magnitudes are very small, indicating that these two channels do not play important roles in our setting. Meanwhile, our main findings on the import tariffs remain robust to the addition of these controls, lending support to the argument of import competition.

As an additional check, we also investigate whether cities that experienced greater reductions in import tariffs also witnessed a larger increase in imports. Because many cities in our dataset report zero import values, to deal with potential estimation bias or sample selection bias we use the Poisson pseudo maximum likelihood estimation devised by Silva and Tenreyro (2006). Specifically, we regress the level of imports on our regressor of interest (i.e., $Tariff_{ct}$) along with a set of city and year dummies and other time-varying controls using the Poisson estimation. The regression results are reported in column 4 of Table 3.3. We find that imports increased in cities experiencing more tariff reduction, which supports the import competition argument

3.4.2 Intra-vs. Inter-Industry Reallocation

The intensification of import competition may lead to a decline of SOE share within each industry (intra-industry reallocation) or a shrinkage of industries that are dominated by SOEs (inter-industry reallocation). Both effects would cause a decline of SOE share at the city level. In other words, our outcome variable, the change in output share of SOEs in city c at time t (y_{ct}), can be decomposed as

$$\begin{aligned}\Delta y_{ct} &= \sum_i \left(\Delta \frac{Ouput_{ict}^{SOE}}{\sum_i Output_{ict}} \right) = \sum_i \left(\Delta s_{ict} \omega_{ict} \right) \\ &= \underbrace{\sum_i \frac{\omega_{ict} + \omega_{ict-1}}{2} \Delta s_{ict}^{SOE}}_{intra-industry} + \underbrace{\sum_i \frac{s_{ict}^{SOE} + s_{ict-1}^{SOE}}{2} \Delta \omega_{ict}}_{inter-industry}\end{aligned} \tag{3.6}$$

where i denotes four-digit industry; $s_{ict}^{SOE} \equiv \frac{Output_{ict}^{SOE}}{Output_{ict}}$ captures the SOE output share in industry i of city c at time t; and $\omega_{ict} \equiv \frac{Output_{ict}}{\sum_i Output_{ict}}$ represents the share of industry i in city c at time t.

Hence, $\Delta y_{ct}^{\text{intra}} = \sum_i \frac{\omega_{ict} + \omega_{ict-1}}{2} \Delta s_{ict}^{SOE}$ captures the resource reallocation from SOEs to non-SOEs within an industry; and $\Delta y_{ct}^{\text{inter}} = \sum_i \frac{s_{ict}^{SOE} + s_{ict-1}^{SOE}}{2} \Delta \omega_{ict}$ captures the resource reallocation from one industry to another. To disentangle the intra- and inter-industry effects of trade liberalization, we conduct two more regressions using $\Delta y_{ct}^{\text{intra}}$ and $\Delta y_{ct}^{\text{inter}}$ as the respective outcome variables.

The regression results are reported in Table 3.4. We find that trade liberalization upon the WTO accession has both negative and statistically significant effects on intra-industry (in column 1) and inter-industry (in column 2) resource reallocation, but the former has a much bigger magnitude than the latter. These results imply that the decline in the SOE output share detected previously is mainly driven by the decline of SOE output share within each industry, whereas across industries there is some evidence that the output of industries with a strong SOE presence prior to China's WTO accession decreased after that.

3.4.3 Extensive-vs. Intensive Margins

We use the growth accounting method to examine the extensive and intensive margin effects of trade liberalization on the output share of SOEs. Specifically, we make two changes to the output share measurement in the baseline Equation (3.3): first, we use the ratio of total SOE output over the total non-SOE output in city c at year t (in logarithm) as the outcome variable (i.e., $\log \frac{Q_{c,t}^{SOE}}{Q_{c,t}^{nonSOE}}$); and second, we look at the change over two time points, 2001 (one year before the WTO accession) and 2005 (four years after the WTO accession and the last year in our sample period), to capture the overall effect:

$$\Delta_{t=2001,2005} \log \frac{Q_{c,t}^{SOE}}{Q_{c,t}^{nonSOE}} = \frac{\Delta_{t=2001,2005} Q_{c,t}^{SOE}}{Q_{c,2001}^{SOE}} - \frac{\Delta_{t=2001,2005} Q_{c,t}^{nonSOE}}{Q_{c,2001}^{nonSOE}}. \tag{3.7}$$

Meanwhile, for each group (i.e., SOEs and non-SOEs), the five-year change in total output can be further decomposed into two parts: change in output of surviving firms and change in output due to entry and exit, i.e.,

$$\Delta_{t=2001,2005} Q_{c,t}^{j} = \Delta_{t=2001,2005} Q_{c,t}^{j,surviving} + \left(Q_{c,2005}^{j,entry} - Q_{c,2001}^{j,exit}\right), \tag{3.8}$$

where $j \in \{SOE, nonSOE\}$.

Combining Equations (3.7)-(3.8), we have

$$\begin{aligned} \Delta_{t=2001,2005} \ln \frac{Q_{c,t}^{SOE}}{Q_{c,t}^{nonSOE}} = & \underbrace{\frac{\Delta_{t=2001,2005} Q_{c,t}^{SOE,surviving}}{Q_{c,2001}^{SOE}} - \frac{\Delta_{t=2001,2005} Q_{c,t}^{nonSOE,surviving}}{Q_{c,2001}^{nonSOE}}}_{Intensive\ Margin} \\ & + \underbrace{\frac{Q_{c,2005}^{SOE,entry} - Q_{c,2001}^{SOE,exit}}{Q_{c,2001}^{SOE}} - \frac{Q_{c,2005}^{nonSOE,entry} - Q_{c,2001}^{nonSOE,exit}}{Q_{c,2001}^{nonSOE}}}_{Extensive\ Margin}. \end{aligned} \tag{3.9}$$

Table 3.5 shows that from 2001 to 2005, the total output of SOEs increased 2.51% from 1,514 billion RMB to 1,552 billion RMB, whereas the total output of non-SOEs increased 148.65% from 6,676 billion RMB to 16,600 billion RMB. This shows that the relative importance of SOEs declined in the post-WTO accession period and is consistent with our empirical findings.

More interestingly, the total output of surviving SOEs actually increased by 230 billion RMB from 2001 to 2005, contributing to 605.3% of the output gain of all SOEs that existed in 2001. However, because there were more exits than entries of SOEs in these five years, the extensive margin effect led to a gross output loss of 191 billion RMB, which implies that the net SOE output gain is only less than 40 billion RMB. Correspondingly, for non-SOEs, the intensive and extensive margins contribute 42.1% and 57.9% to their net output gain respectively. These numbers suggest that the decline in SOE output share is driven by entry and exit (or the extensive margin effect) instead of a substantial contraction of output among surviving firms (or the intensive margin effect).

To further investigate the effect of trade liberalization on the external and internal margins of SOE output share, we run two regressions similar to the regression in Equation (3.3). First, we specify the intensive margin effect by

$$\Delta \ln \frac{Q_{c,t}^{SOE,surviving}}{Q_{c,t}^{nonSOE,surviving}} = \beta \Delta Tariff_{ct} + \Delta X_{ct}^{'}\gamma + \Delta \left(Z_{c2001} \cdot Post02_{t}\right)^{'} \cdot \theta + \lambda_{pt} + \Delta\varepsilon_{ct,} \quad (3.10)$$

where $t \in \{2001,2005\}$. Next, the specification for the extensive margin effect is given by

$$\Delta m_{ct} = \beta \Delta Tariff_{ct} + \Delta X_{ct}^{'}\gamma + \Delta \left(Z_{c2001} \cdot Post02_{t}\right)^{'} \cdot \theta + \lambda_{pt} + \Delta\varepsilon_{ct,} \quad (3.11)$$

where $\Delta m_{ct} = \frac{Q_{c,2005}^{SOE,entry} - Q_{c,2001}^{SOE,exit}}{Q_{c,2001}^{SOE}} - \frac{Q_{c,2005}^{nonSOE,entry} - Q_{c,2001}^{nonSOE,exit}}{Q_{c,2001}^{nonSOE}}$.

The regression results are reported in Table 3.6. Column 1 shows that the intensive margin effect of trade liberalization on the SOE output share is highly insignificant, while the extensive margin effect (column 2) is significant and economically meaningful.

In summary, we find that the decline in the SOE output share after China's WTO accession is primarily caused by the exit of SOEs instead of a reduction in the output of surviving SOEs. These results are in line with Brandt, Van Biesebroeck, Wang, and Zhang (2012), who also look at the episode of China's WTO accession and find that much of the sectoral productivity gains take place at the extensive margin.

3.4.4 Heterogeneous Response of SOEs

Building on the finding that much of the trade liberalization effect happens at the extensive margin, we further investigate if there are any common characteristics among the SOEs that exited the market.

Productivity difference. We start with the differential exit rates of SOEs at different productivity levels. According to the firm heterogeneity literature (e.g., Melitz 2003), trade liberalization drives out the least productive firms as import competition raises the survival threshold. To test this hypothesis, we first divide SOEs into four quantiles (i.e., top 25%, 25%-50%, 50%-75%, and bottom 25%) based on their productivity levels in 2001, which are estimated using the method devised in Olley and Pakes (1996). Next, we calculate the exit rate from 2001 to 2005 for each quantile and each city:

$Exit_{q,c} \equiv \frac{\Delta_{t=2001:2005} N_{q,c,t}}{N_{q,c,2001}}$, where q denotes the productivity quantile in 2001.

The regression specification is

$$Exit_{q,c} = \beta_q \Delta Tariff_{ct} + \Delta X_{ct}^{'} \gamma + \Delta\left(Z_{c2001} \cdot Post02_t\right)^{'} \cdot \theta + \lambda_{pt} + \Delta\varepsilon_{ct}, \qquad (3.12)$$

where β_q is the coefficients for different quantiles. To further capture the city-specific differential time trend between 2001 and 2005, we include two additional controls: the entry rate for SOEs and the exit rate for non-SOEs.

The regression results are shown in column 1 of Table 3.7. We find that all four estimated coefficients are positive, suggesting that trade liberalization increases the exit rate. However, only one of the estimates is statistically significant. In terms of magnitude, we find that the coefficient estimate for the top 25% category is smaller than the one for the 25%-50% category, which is in turn smaller than the coefficient for the 50%-75% category. These are consistent with the firm heterogeneity literature's argument that less productive firms are more likely to exit the market after trade liberalization.

Surprisingly, however, the estimated coefficient for the bottom 25% category is found to be the smallest, suggesting that the weakest SOEs have the highest probability to survive increased import competition upon the WTO accession. One plausible explanation is that the weakest SOEs (in terms of productivity) continued to enjoy government protection in the post-WTO accession period and thus were largely shielded from the increasingly fierce import competition. To explore this possibility, we further investigate the differential exit rates among SOEs affiliated with different layers of the Chinese government.

Affiliation difference. In China, different SOEs are administered by different layers of the government: the central government, provinces, cities, counties, and townships. One expects that SOEs administered by higher-level governments would enjoy more protection than those administered by lower-level governments for at least two reasons: (a) higher-level governments have more decision power given the top-down nature of Chinese politics; (b) lower-level governments in China have been encountering structural fiscal problems since the 1994 fiscal reform (World Bank and Development Research Center of the State Council, P.R.C. 2012). As such, we use the affiliation of an SOE as a proxy for the degree of government protection it received and investigate whether SOEs administrated by different levels of the Chinese government responded to WTO accession differently. Based on the information of SOE affiliation in 2001, we classified the SOEs into three categories: SOEs under the administration of above-city-level governments (i.e., central and provincial governments), SOEs under the administration of the city government, and SOEs under the administration of below-city-level governments (i.e., county and township governments). We then calculate the exit rate of SOEs for each of these three categories and each city from 2001 to 2005: $Exit_{a,c} \equiv \frac{\Delta_{t=2001:2005} N_{a,c,t}}{N_{q,c,2001}}$, where a denotes the SOE affiliation category. The regression specification is

$$Exit_{a,c} = \beta_a \Delta Tariff_{ct} + \Delta X_{ct}^{'} \gamma + \Delta \left(Z_{c2001} \cdot Post02_t \right)^{'} \cdot \theta + \lambda_{pt} + \Delta \varepsilon_{ct,} \quad (3.13)$$

where β_a is the coefficient for category a. To further capture the city-specific differential time trend between 2001 and 2005, we include two additional controls, the entry rate for SOEs and the exit rate for non-SOEs.

Regression results are reported in column 2 of Table 3.7. We find that the estimated coefficient on SOEs affiliated with below-city-level administration is positive and statistically significant, implying that trade liberalization induced the exit of SOEs in this category. Meanwhile, the estimated coefficients on SOEs in the other two categories are negative and insignificant, indicating that these SOEs were barely affected by trade liberalization.

How do we reconcile the results in columns 1–2 of Table 3.7? Table 3.8 provides the breakdown of SOE productivity quantile (as defined in column 1) by affiliation (as defined in column 2). Interestingly, we find that the weakest SOEs (in terms of productivity in 2001) were most likely to be administered by the central, provincial, or city governments.

In sum, these results suggest that the aberration in the differential exit rates across productivity quantiles is driven by differential levels of government protection. They also indicate that although trade liberalization led to the reallocation of resources from inefficient SOEs to efficient non-SOEs, it does not completely eliminate the existing distortions and some (perhaps substantial) deadweight loss continues to exist in the post-WTO accession period.

3.5 Conclusion

In this paper, we study how China's accession to the WTO impacts the market share of inefficient but politically favored SOEs. We find that tariff reductions following China's WTO accession led to a decline in the SOE output share. This result is robust to a variety of robustness checks. In our preferred specification, China's WTO accession led to a 2.5 percentage-point decline in the SOE output share, which corresponds to a 1.4% reduction in the standard deviation of manufacturing firm productivity in China. Given the general lack of competitiveness of Chinese SOEs, we interpret the reduction in their output share and the subsequent narrowing of productivity dispersion among the remaining firms as an improvement of allocative efficiency.

We further verify that the post-WTO accession contraction of SOE output share was driven by increased import competition instead of improved access to overseas markets or cheaper imported intermediate goods. The SOE output share decline was broad-based and not limited to selected industries where the SOEs were dominant.

Importantly, we find that the SOE output share decline took place at the extensive margin, and it was the exit of SOEs affiliated with the lowest levels of government (counties and townships) that drove the decline. By contrast, SOEs affiliated with the central, provincial, and city governments were barely affected even though many of the least productive manufacturing firms in China belonged to this group. This shows that while trade discourages political favoritism and improves resource allocation, the welfare gains are made only at the margin, and some inefficiency is likely to persist as long as the government has the financial ability to pay for its support of inefficient enterprises. In other words, trade mitigates but does not solve the problem. Another way of interpreting this finding is to think of trade as a catalyst that lessens existing political distortion of resource allocation. The effect of this catalyst strengthens as the government's ability to provide discriminatory treatment weakens.

Our findings contribute to the literature by identifying another channel (the improvement in resource allocation in the presence of political economy distortions) through which trade benefits a nation. With regard to China, now the world's biggest trading nation, we find that different layers of the Chinese government responded differently to WTO accession: some increased their support to the SOEs under their watch, while others withdrew theirs. Given the central and complex role that the state plays in China's economic development, our findings suggest that it is useful and perhaps even important to treat the Chinese state as an agglomeration of component parts instead of a unitary government when studying its behaviors and decisions in international trade.

Figures and Tables for Chapter One

Figure 1.1: The Ratio of Exports in Our Regression Sample over Total Exports (2000-2006)

Figure 1.2: Monthly Nominal RMB Exchange Rate (2000-2006)

Figure 1.3: Difference between Exports to U.S. and Exports to Other Non-U.S. Countries in Our Regression Sample

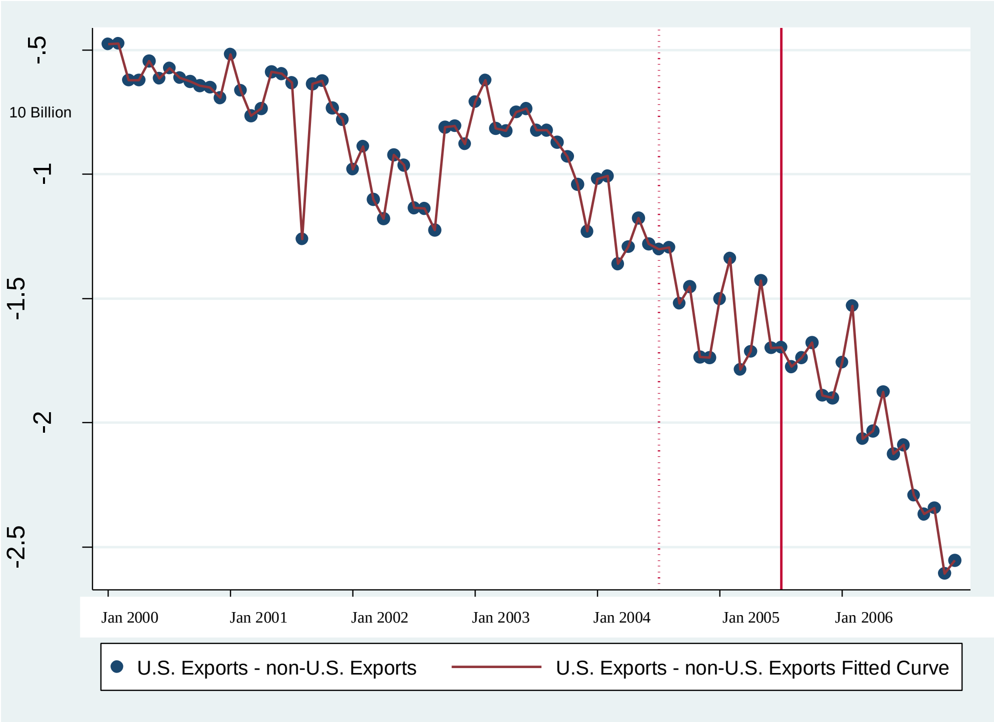

Table 1.1: List of Countries

Australia	Finland	Malta	Russian Federation
Austria	France	Mauritius	Singapore
Belgium	Germany	Mexico	Slovak Republic
Brazil	Greece	Morocco	Slovenia
Bulgaria	Hungary	Myanmar	South Africa
Cameroon	Iceland	Nepal	Spain
Canada	Indonesia	Netherlands	Sri Lanka
Chile	Ireland	New Zealand	Sweden
Colombia	Israel	Norway	Switzerland
Costa Rica	Italy	Papua New Guinea	Thailand
Croatia	Japan	Paraguay	Turkey
Czech Republic	Korea, Republic of	Poland	United Kingdom
Denmark	Luxembourg	Portugal	United States
Estonia	Madagascar	Romania	Uruguay

Table 1.2: Main Results

	(1)	(2)	(3)	(4)
Dependent Variable	Ln(Export Value)			
Treatment*Post	-0.176***	-0.154***	-0.154***	-0.165***
	(0.052)	(0.054)	(0.054)	(0.059)
Time Fixed Effect	X	X	X	X
Country Fixed Effect	X	X	X	
Ln (Import Value)		X	X	X
Producer Heterogeneity			X	X
Country-Specific Month-of-Year Effect				X
Number of Observations	4704	4704	4704	4704
R-squared	0.0883	0.2825	0.2829	0.3515

Notes: Standard errors, clustered at the country level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 1.3: Robustness Checks

	(1)	(2)	(3)	(4)	(5)
Dependent Variable	Ln (Export Value)				
Specification	Incl. Country Time Trend	Incl. Pre-Reform Trend	2004-2006	Homogeneous	Triple Difference
Treatment*Post	-0.068**		-0.087**	0.037	
	(0.032)		(0.037)	(0.080)	
Treatment*08/2005 onward		-0.088**			
		(0.043)			
Treatment*07/2004-06/2005		-0.006			
		(0.025)			
Treatment*Before 07/2004		0.095***			
		(0.035)			
Treatment*Post*Differentiated					-0.250***
					(0.077)
Time Fixed Effect	X	X	X	X	
Country Fixed Effect	X	X	X	X	
Ln (Import Value)	X	X	X	X	X
Producer Heterogeneity	X	X	X	X	X
Country-Specific Month-of-Year Effect	X	X	X	X	
Country-Time Fixed Effect					X
Country-Product Fixed Effect					X
Product-Time Fixed Effect					X
Number of Observations	4704	4704	2016	3609	7218
R-squared	0.2725	0.3515	0.2486	0.2124	0.0171

Notes: Standard errors, clustered at the country level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 1.4: Exchange Rate Elasticity

	(1)	(2)	(3)	(4)
Specification	OLS		2SLS	
	Nominal Exchange Rate	Real Exchange Rate	Nominal Exchange Rate	Real Exchange Rate
Dependent Variable	Ln (Export Value)			
Ln (Exchange Rate)	-0.454**	-0.685**	-1.605***	-1.125***
	(0.190)	(0.332)	(0.546)	(0.359)
F test of Excluded Instruments			[27.02]	[59.21]
Time Fixed Effect	X	X	X	X
Ln (Import Value)	X	X	X	X
Producer Heterogeneity	X	X	X	X
Country-Specific Month-of-Year Effect	X	X	X	X
Number of Observations	4704	4367	4704	4367
R-squared	0.3460	0.3508	0.3522	0.328

Notes: Standard errors, clustered at the country level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 1.5: Trade Diversion

	(1)	(2)
Dependent Variable	Ln (Export Value)	
Specification	Exclude OECD	OCED versus the Rest (excl. U.S.)
Treatment*Post	-0.186**	-0.029
	(0.091)	(0.116)
Time Fixed Effect	X	X
Ln (Import Value)	X	X
Producer Heterogeneity	X	X
Country-Specific Month-of-Year Effect	X	X
Number of Observations	2268	4620
R-squared	0.3787	0.3544

Notes: Standard errors, clustered at the country level, are reported in the parenthesis.
***p<0.01, ** p<0.05, * p<0.1.

Table 1.6: The Effect of Exchange Rate Reform on Extensive and Intensive Margins

	(1)	(2)	(3)	(4)	(5)
	Extensive Margin		Intensive Margin		
Dependent Variable	Ln(Number of Firms)	Ln(Number of HS8)	Ln(Price)	Ln(Quantity)	Ln(Revenue)
Treatment*Post	-0.066**	-0.292***	-0.013***	-0.027	-0.041***
	(0.029)	(0.029)	(0.005)	(0.020)	(0.018)
Time Fixed Effect	X	X	X	X	X
Product Fixed Effect	X	X	X	X	X
Ln(Import Value)	X	X	X	X	X
Producer Heterogeneity	X	X	X	X	X
Country-Specific Month-of-Year Effect	X	X	X	X	X
Number of Observations	4704	4704	3522562	3522562	3522562

Notes: Standard errors, clustered at the country level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 1.7: Heterogeneous Effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Dependent Variable	Ln (Export Value)								
Specification	Coastal vs. Inland	SOE vs. Private	Time Sensitive vs. Insensitive	Consumer Food Products	Consumer Non-food Non-durables	Consumer Durables	Intermediate Goods	Energy	Capital Goods
Treatment*Post*Group	-0.432***	-0.187***	-0.134**						
	(0.056)	(0.040)	(0.057)						
Treatment*Post				-0.028	-0.176***	-0.280***	-0.107*	0.139	-0.363***
				(0.082)	(0.069)	(0.080)	(0.057)	(0.193)	(0.127)
Country-Time Fixed Effect	X	X	X						
Country-Group Fixed Effect	X	X	X						
Group-Time Fixed Effect	X	X	X						
Time Fixed Effect				X	X	X	X	X	X
Ln (Import Value)	X	X	X	X	X	X	X	X	X
Producer Heterogeneity	X	X	X	X	X	X	X	X	X
Country-Specific Month-of-Year Effect				X	X	X	X	X	X
Number of Observations	5880	3192	5712	4444	4704	4700	4704	2582	4696
R-squared	0.0264	0.9112	0.0882	0.7085	0.3854	0.5356	0.3291	1.6111	0.8361

Notes: Standard errors, clustered at the country level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Figures and Tables for Chapter Two

Figure 2.1: Elasticity of Substitution and PRODY (lowess line)

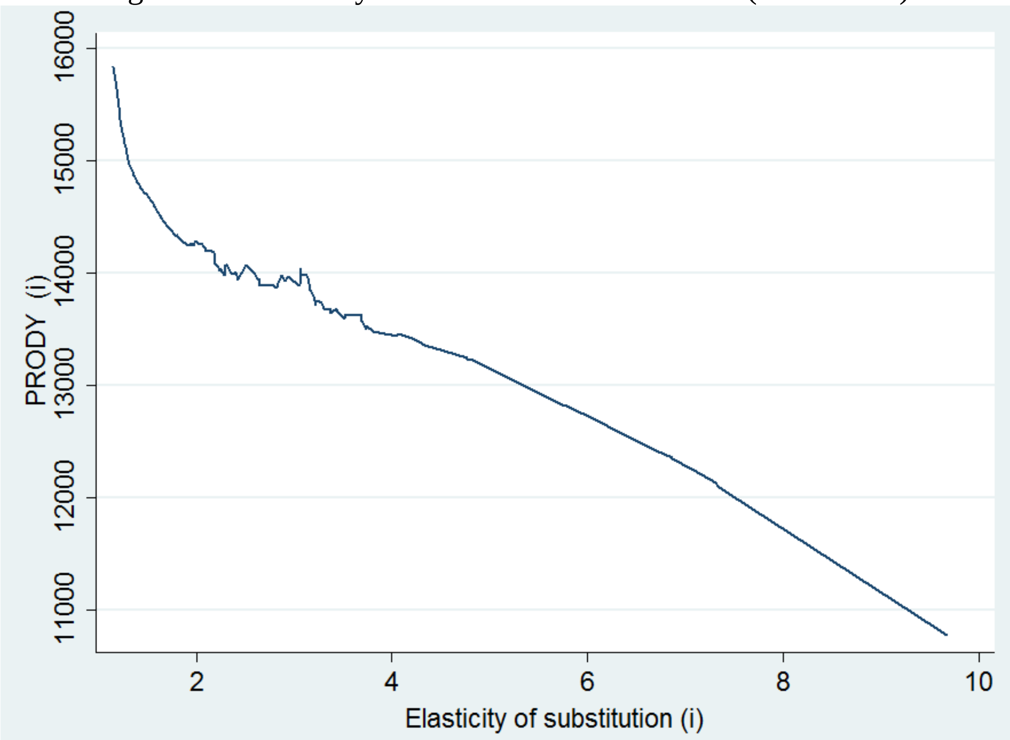

Figure 2.2: Monthly Nominal USD/RMB (2000-2006)

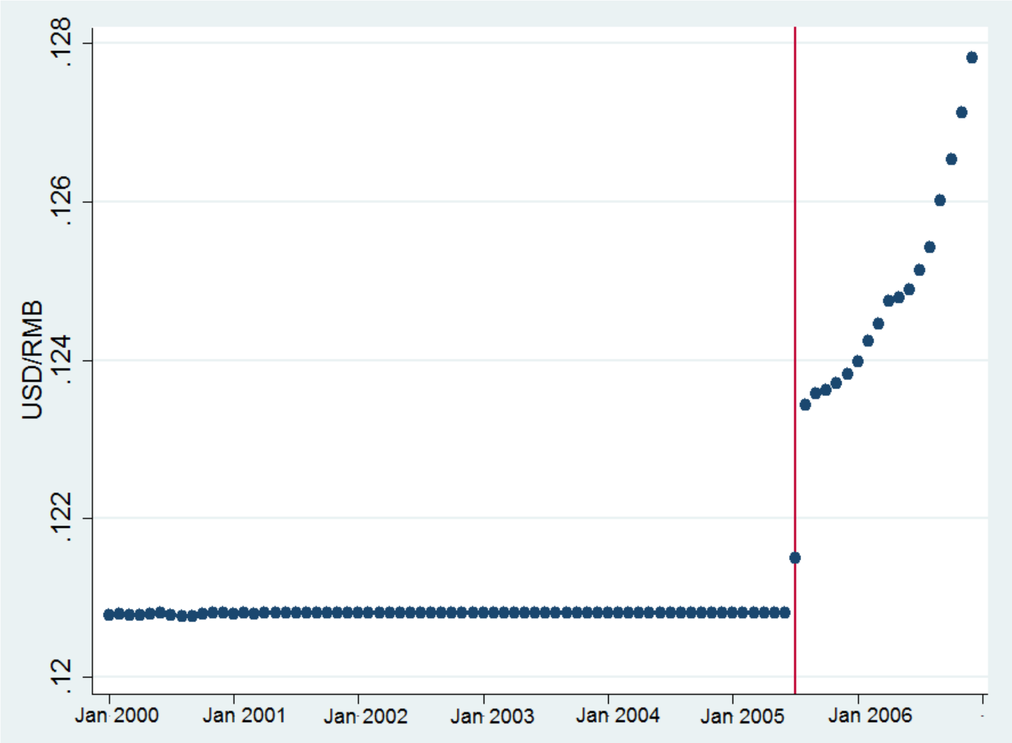

Figure 2.3: Ln(Export Similarity Index) 2003 versus 2005

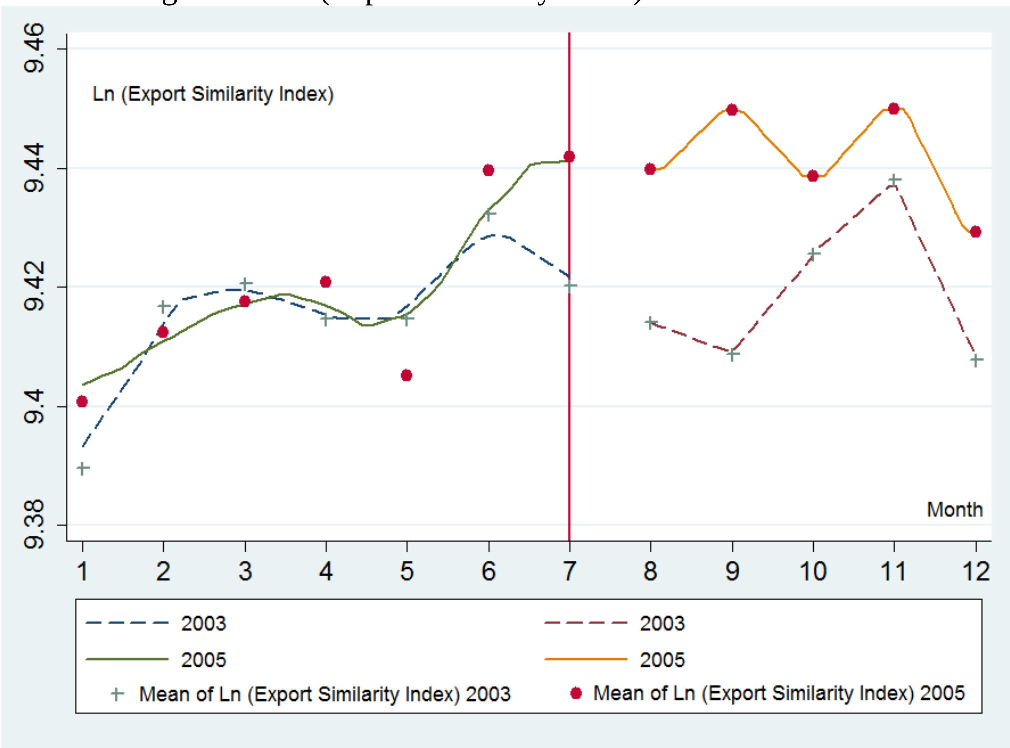

Table 2.1: The Top 5 Sophisticated Goods and the Bottom 5 Sophisticated Goods (U.S.$2000)

	Product name	PRODY 2000
Bottom 5	Vegetable products nes	739.67145
	Asses, mules, and hinnies, live	803.94128
	Sisal and agave, raw	822.37665
	Cloves (whole fruit, cloves, and stems)	866.57587
	Hand-made lace, in the piece, in strips or in motifs	901.80627
Top 5	Flat-rolled iron or non-alloy steel, coated with aluminium, width > 600mm	50699.391
	Sheet piling of iron or steel	46986.039
	Sections H iron or non-alloy steel, nfw hot-roll/drawn/extruded > 80m	46242.609
	Tyre cord fabric of viscose rayon	46077.578
	Cermets and articles thereof, waste or scrap	46058.699

Table 2.2: Main Results

	1	2
	RD	DD
Dependent Variable	Ln (Export Similarity Index)	
August	-0.002	
	(0.023)	
August* Year 2005		0.018**
		(0.006)
Month-of-Year Fixed Effect		X
Year Fixed Effect		X
Number of Observations	4404	8525
R-squared		0.0020

Notes: Standard errors in Column 1 are bootstrapped; standard errors in Column 2 are clustered at the month level. Standard errors are reported in the parentheses. ** p<0.05.

Table 2.3: Robustness Checks

	1 Nigeria DD	2 DDD	3 Placebo Test, 2003 vs 2002	4 U.S. Exports to China	5 Exclusion of Processing Trade
Dependent Variable			Ln (Export Similarity Index)		
August*Year2005	0.001	0.017	-0.001	-0.049**	0.014*
	(0.015)	(0.011)	(0.009)	(0.022)	(0.008)
Month-of-Year Fixed Effect	X		X	X	X
Year Fixed Effect	X		X	X	X
Group-Month-of-Year Fixed Effect		X			
Group-year Fixed Effect		X			
Month-year Fixed Effect		X			
Number of Observations	3972	12497	8262	7015	8435
R-squared	0.0074	0.0074	0.0015	0.0123	0.0017

Notes: Standard errors in Column 1, 3, 4, 5, are clustered at the month level; standard errors in Column 2 are clustered at country-month level; standard errors are reported in the parentheses. ** p<0.05, * p<0.1.

Table 2.4: Decomposition of the Effect of Currency Appreciation

	1	2
	Across Firms within a City	Across Products within a Firm
Dependent Variable		Ln (Export Similarity Index)
August* Year 2005	0.015***	0.002***
	(0.005)	(0.001)
Month-of-Year Fixed Effect	X	X
Year Fixed Effect	X	X
City Fixed Effect	X	
Firm Fixed Effect		X
Number of Observations	8516	120672
R-squared	0.6311	0.8827

Notes: Standard errors are clustered at the month level. Standard errors are reported in the parentheses. *** p<0.01.

Figures and Tables for Chapter Three

Figure 3.1: Tariffs (1996-2007)

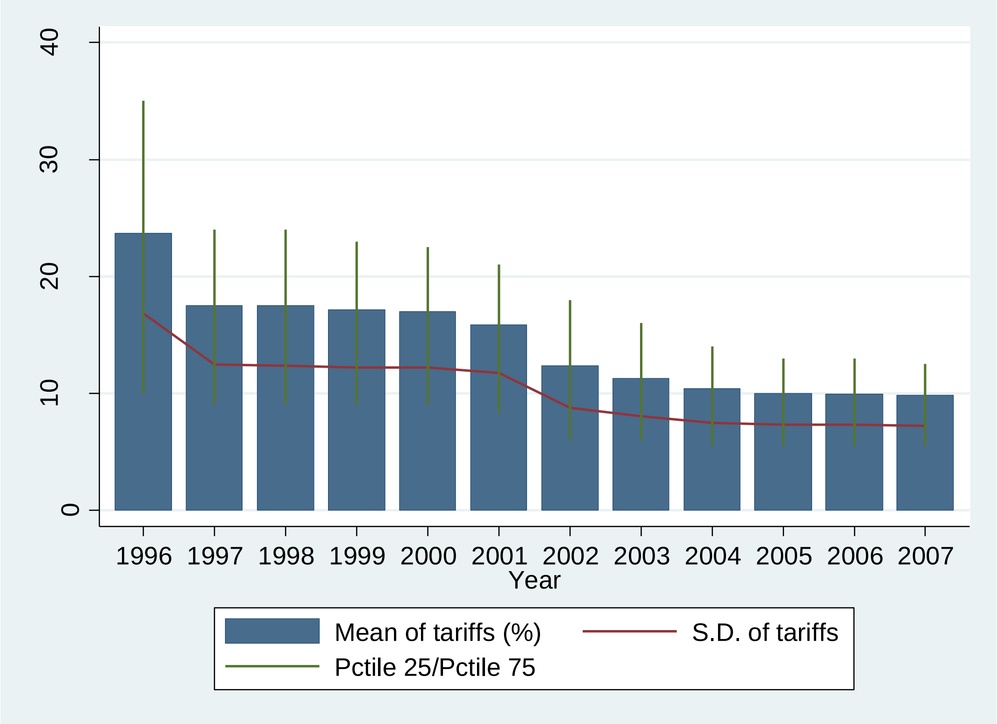

Figure 3.2: The Correlation between Tariffs in 2001 and Tariff Changes during 2001-2005

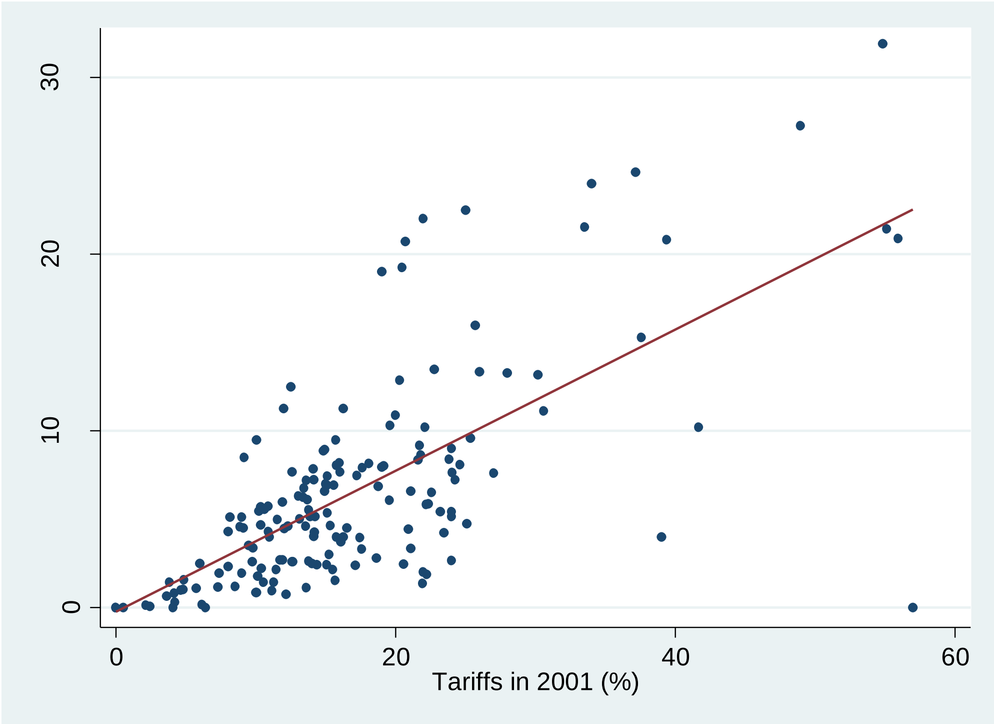

Figure 3.3: Estimated Coefficients from the Flexible DID Estimation

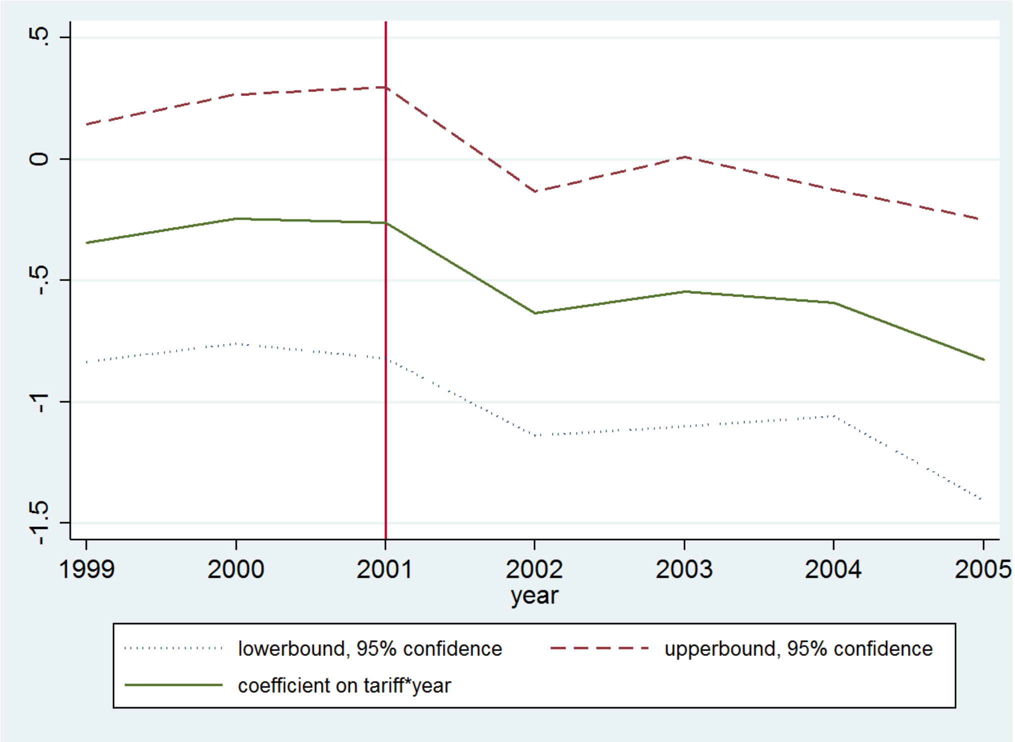

Table 3.1: Main Results

	(1)	(2)
Dependent Variable	Δ Output Share of SOEs	
ΔTariff*Post02	-0.254**	-0.304***
	(0.125)	(0.107)
Province-Year Fixed Effects	X	X
Δ Ln(GDP per capita)	X	X
Δ Gov Consumption/GDP	X	X
Δ Ln(Dairy per Capita)*Post02		X
Δ Ln(Vegetable per Capita)*Post02		X
Δ Ln(Telephone)*Post02		X
Δ University*Post02		X
Δ Coastal*Post02		X
Δ Northern*Post02		X
Δ Mountain*Post02		X
Number of Observations	1813	1697
R-squared	0.2184	0.2293

Notes: Standard errors, clustered at the city level, are reported in the parenthesis. *** p<0.01, ** p<0.05.

Table 3.2: Robustness Checks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable		Δ Output Share of SOEs			Δ Privatization	Δ Output Share of SOEs	
Specification	Incl. City-linear trend	Incl. FDI	Incl. Privatization	Continuous SOE and non-SOE		Pre-WTO	Pure Exporters
ΔTariff*Post02	-0.271**	-0.295***	-0.306***	-0.293**	-0.030		0.189
	(0.136)	(0.110)	(0.107)	(0.146)	(0.062)		(0.359)
ΔTariff						-0.087	
						(0.297)	
Province-Year Fixed Effects	X	X	X	X	X	X	X
Δ Ln(GDP per capita)	X	X	X	X	X	X	X
Δ Gov Consumption/GDP	X	X	X	X	X	X	X
Δ Ln(Dairy per Capita)*Post02	X	X	X	X	X		X
Δ Ln(Vegetable per Capita)*Post02	X	X	X	X	X		X
Δ Ln(Telephone)*Post02	X	X	X	X	X		X
Δ University*Post02	X	X	X	X	X		X
Δ Coastal*Post02	X	X	X	X	X		X
Δ Northern*Post02	X	X	X	X	X		X
Δ Mountain*Post02	X	X	X	X	X		X
City-Specific Linear Trend	X						
Δ Ln(FDI)		X					
Δ Privatization			X				
Number of Observations	1697	1621	1482	1697	1482	721	1040
R-squared	0.2761	0.2444	0.1994	0.2579	0.2591	0.3287	0.4182

Notes: Standard errors, clustered at the city level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.3: Imports, Exports, and Imported Inputs

	(1)	(2)	(3)	(4)
Dependent Variable	Δ Output Share of SOEs			Imports
ΔTariff*Post02	-0.344***	-0.240*	-0.283**	0.991*
	(0.113)	(0.125)	(0.124)	(0.563)
ΔExport Tariff*Post02	0.002		0.002	
	(0.001)		(0.001)	
ΔInput Tariff*Post02		-0.094	-0.094	
		(0.129)	(0.137)	
Province-Year Fixed Effects	X	X	X	X
Δ Ln(GDP per capita)	X	X	X	X
Δ Gov Consumption/GDP	X	X	X	X
Δ Ln(Dairy per Capita)*Post02	X	X	X	X
Δ Ln(Vegetable per Capita)*Post02	X	X	X	X
Δ Ln(Telephone)*Post02	X	X	X	X
Δ University*Post02	X	X	X	X
Δ Coastal*Post02	X	X	X	X
Δ Northern*Post02	X	X	X	X
Δ Mountain*Post02	X	X	X	X
Number of Observations	1609	1697	1626	1224
R-squared	0.2374	0.2295	0.2376	
Pseudo R2				0.9932

Notes: Standard errors, clustered at the city level, are reported in the parenthesis.
*** p<0.01, ** p<0.05, * p<0.1.

Table 3.4: Intra- vs. Inter-Industry Reallocation

	(1) Intra-Industry Reallocation	(2) Inter-Industry Reallocation
Δ Tariff*Post02	-0.355***	-0.038***
	(0.102)	(0.013)
Province-Year Fixed Effects	X	X
Δ Ln(GDP per capita)	X	X
Δ Gov Consumption/GDP	X	X
Δ Ln(Dairy per Capita)*Post02	X	X
Δ Ln(Vegetable per Capita)*Post02	X	X
Δ Ln(Telephone)*Post02	X	X
Δ University*Post02	X	X
Δ Coastal*Post02	X	X
Δ Northern*Post02	X	X
Δ Mountain*Post02	X	X
Number of Observations	1697	1697
R-squared	0.2476	0.2293

Notes: Standard errors, clustered at the city level, are reported in the parenthesis.
*** p<0.01, ** p<0.05.

Table 3.5: Intensive vs. Extensive Margins, Overall

	(1)	(2)	(3)	(4)	(5)	(6)
	Total		Intensive Margin		Extensive Margin	
Unit (billion RMB)	SOEs	non-SOEs	SOEs	non-SOEs	SOEs	non-SOEs
2001	1,514	6,676	1,023	4,621	491	2,056
2005	1,552	16,600	1,253	8,798	300	7,802
Δ	38	9,924	230	4,177	-191	5,746
	2.51%	148.65%	22.48%	90.39%	-38.90%	279.47%
Intensive/Extensive Margins			605.26%	42.09%	-502.63%	57.90%

Table 3.6: Intensive vs. Extensive Margins, Regression Results

	(1) Intensive Margin	(2) Extensive Margin
ΔTariff*Post02	-0.179	-5.392**
	(1.758)	(2.618)
Province-Year Fixed Effect	X	X
Δ Ln(GDP per capita)	X	X
Δ Gov Consumption/GDP	X	X
Δ Ln(Dairy per Capita)*Post02	X	X
Δ Ln(Vegetable per Capita)*Post02	X	X
Δ Ln(Telephone)*Post02	X	X
Δ University*Post02	X	X
Δ Coastal*Post02	X	X
Δ Northern*Post02	X	X
Δ Mountain*Post02	X	X
Number of Observations	248	240
R-squared	0.2694	0.2162

Notes: Robust standard errors are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.7: Differential Exit Rates

	(1)	(2)
Dependent Variable	SOE exit rate	
Tariff(2001)*TFP Top 25%	0.140	
	(0.233)	
Tariff(2001)*TFP 25%-50%	0.298	
	(0.231)	
Tariff(2001)*TFP 50%-75%	0.533***	
	(0.198)	
Tariff(2001)*TFP Bottom 25%	0.042	
	(0.194)	
Tariff(2001)*Below city-level		0.657***
		(0.219)
Tariff(2001)*city-level		-0.285
		(0.175)
Tariff(2001)*Above city-level		-0.177
		(0.134)
Group dummies	X	X
Province fixed effects	X	X
ln(GDP per capita)	X	X
gov consumption/gdp	X	X
ln(dairy)	X	X
ln(vegetable per capita)	X	X
ln(telephone)	X	X
university	X	X
coast	X	X
north	X	X
mountain	X	X
non-SOE exit rate	X	X
SOE entry rate	X	X
Number of Observations	1003	712
R-squared	0.2291	0.4152

Notes: Robust standard errors are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.8: Breakdown of SOE Productivity Quantile by Affiliation

TFP Quantile	Above city-level SOEs share	City-level SOEs share	Below city-level SOEs share
TFP Top 25%	0.22	0.31	0.47
TFP 25%-50%	0.23	0.27	0.50
TFP 50%-75%	0.28	0.26	0.46
TFP Bottom 25%	0.36	0.27	0.38

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Appendices

Appendix 1 for Chapter One

In this Appendix, we outline a partial equilibrium model to illustrate how an exogenous shock to exchange rate affects exporting behavior. Specifically, we extend Melitz and Ottaviano (2008)'s model to incorporate the role of exchange rate movement.[51] There are totally $N+1$ countries, a Home country (H) and N foreign countries, indexed by $i \in \{1,\ldots,N\}$. Each firm produces a unique variety, competes in the monopolistic-competition manner, and is indexed by its productivity level φ that is drawn from a cumulative distribution function $G(\varphi)$. Without loss of generality, we look only at how the change in Home country's exchange rate against foreign country i affects its exports to that foreign country.

The inverse demand function for a variety produced by firm φ from Home and exported to foreign country i is:[52]

$$p_i(\varphi)e_i = \alpha - \gamma q_i(\varphi) - \eta Q_i, \tag{1.6}$$

where $p_i(\varphi)$ are FOB export prices in foreign country i denominated in Home currency, respectively; e_i is the exchange rate of Foreign currency against Home currency (hence, an increase in e_i means an appreciation in Home currency against foreign country $i's$); $q_i(\varphi)$ is the demand of variety

[51] Similar results regarding the effect of exchange rate on exports can be derived using another commonly-used model, i.e., Melitz (2003)'s framework. See also Berman, Martin, and Mayer (2012).
[52] This inverse demand function can be derived from the maximization of a quadratic linear utility function. For more details, see Melitz and Ottaviano (2008).

φ in foreign country i; and $Q_i \equiv \int_{\varphi} q_i(\varphi)d\varphi$ is the total demand in foreign country i. The demand parameters $\alpha, \gamma,$ *and* $\eta,$ are all positive.

Profit maximization yields the following equilibrium FOB export price:[53]

$$p_i^*(\varphi) = \frac{1}{2}\omega\tau_i\left(\frac{1}{\varphi^*} + \frac{1}{\varphi}\right), \tag{1.7}$$

where $\frac{1}{\varphi^*} \equiv \frac{1}{\omega e_i}\frac{\alpha - \eta Q_i}{\tau_i}$ is the productivity threshold of exporting, that is, the level for which operating profits from foreign country i are zero; ω is the Home wage rate (denominated in Home currency); and $\tau_i > 1$ is the iceberg trade cost between Home and foreign country i (i.e., for every τ_i units shipped, only one unit arrives at the destination). For an active exporter φ in Home, its export volume to foreign country i is:

$$q_i^*(\varphi) = \frac{1}{2}\omega\tau_i e_i\left(\frac{1}{\varphi^*} - \frac{1}{\varphi}\right). \tag{1.8}$$

Hence, the aggregate export value V_i (denominated in Home currency) from Home to foreign country i is the sum of all active individual exporters' export revenues ($r(\varphi) \equiv p_i^*(\varphi)q_i^*(\varphi)$), i.e.,

$$V_i = \int_{\varphi_i^*}^{\infty} r(\varphi)dG(\varphi) = \int_{\varphi_i^*}^{\infty} p_i^*(\varphi)q_i^*(\varphi)dG(\varphi). \tag{1.9}$$

And the effect of the change in the exchange rate e_i on the aggregate export value V_i is

[53] Here we abuse the term FOB price a little, because $p_i(\varphi)$ includes the trade cost τ_i. In the gravity model of our empirical part, we control for τ_i with country fixed effects.

$$\frac{\partial V_i}{\partial e_i} = \underbrace{\int_{\varphi_i^*}^{\infty} \frac{\partial\left[p_i^*(\varphi)q_i^*(\varphi)\right]}{\partial e_i} dG(\varphi)}_{\text{intensive margin}} - \underbrace{p_i^*(\varphi)q_i^*(\varphi)G'(\varphi_i^*)\frac{\partial \varphi_i^*}{\partial e_i}}_{\text{extensive margin}}.$$

(1.10)

The first term on the right-hand of Equation (1.10) represents the effect from continuing exporters (or the intensive-margin effect), which can be shown to be negative, i.e.,

$$\frac{\partial r(\varphi)}{\partial e_i} = \frac{\partial\left[p_i^*(\varphi)q_i^*(\varphi)\right]}{\partial e_i} < 0 \forall \varphi \geq \varphi_i^*. \tag{1.11}$$

Meanwhile, the intensive-margin effect can be further decomposed into a price effect ($\frac{\partial p_i^*(\varphi)}{\partial e_i}$) and a volume effect ($\frac{\partial q_i^*(\varphi)}{\partial e_i}$), both of which can be proved to be negative, i.e.,

$$\underbrace{\frac{\partial p_i^*(\varphi)}{\partial e_i}}_{\text{price effect}} < 0, \quad \underbrace{\frac{\partial q_i^*(\varphi)}{\partial e_i}}_{\text{quantity effect}} < 0 \quad \forall \varphi \geq \varphi_i^*. \tag{1.12}$$

The second term on the right-hand of Equation (1.10) captures the extensive-margin effect, that is, the effect due to the change in the number of exporters, which is a monotonically decreasing function of φ_i^*. It can be proved that the productivity threshold of exporting φ_i^* is an increasing function of e_i, i.e,

$$\frac{\partial \varphi_i^*(\varphi)}{\partial e_i} > 0, \tag{1.13}$$

therefore, we have a negative extensive margin effect of a currency appreciation.

Combining Equations (1.11) and (1.13), we have

$$\frac{\partial V_i}{\partial e_i} < 0, \tag{1.14}$$

that is, an appreciation in Home currency against foreign country $i's$ results in a decrease in aggregate export value for Home to foreign country i.

The intuition for Equation (1.14) is as follows. There is an incomplete pass-through of an exchange rate appreciation: Home exporters absorb partially the appreciation effect by lowering its FOB export prices, but final prices (denominated in foreign country $i's$ currency) in foreign country i still increase, which consequently leads to a fall in the final demand. As a result, such incomplete pass-through reduces FOB export revenues that Home exporters can obtain in foreign country i, and hence decreases the aggregate export value to that country. Moreover, given that the reduction in export revenue is more significant for less productive exporters (i.e., $\frac{\partial^2 r(\varphi)}{\partial e_i \partial \varphi}$), some (least productive) exporters find it not profitable to sell in and hence choose to exit foreign country i, which further decreases the aggregate export value to that country.

Appendix 2 for Chapter Two

Proof of Proposition 1

Using (1), we can define the equilibrium entry ratio m^* to be the solution to $F(m,e,\sigma_1,\sigma_2,\alpha_1,\alpha_2)=0,$ where

$$\begin{aligned}F(m,e,\sigma_1,\sigma_2,\alpha_1,\alpha_2)&\equiv\frac{\alpha_1}{\sigma_1}\left[\frac{L_n\left(1-e^{-\sigma_1}\tau^{1-\sigma_1}\right)}{m+(e\tau)^{1-\sigma_1}}+\frac{L_s\left(e^{\sigma_1}\tau^{1-\sigma_1}-1\right)}{1+m\left(e^{-1}\tau\right)^{1-\sigma_1}}\right]\\&+\frac{\alpha_2}{\sigma_2}\left[\frac{L_n\left(1-e^{-\sigma_2}\tau^{1-\sigma_2}\right)}{m+(e\tau)^{1-\sigma_2}}+\frac{L_s\left(e^{\sigma_2}\tau^{1-\sigma_2}-1\right)}{1+m\left(e^{-1}\tau\right)^{1-\sigma_2}}\right]\\&\equiv H(m,e,\sigma_1,\alpha_1)+H(m,e,\sigma_2,\alpha_2),\end{aligned}$$

where

$$H(m,e,\sigma,\alpha)\equiv\frac{\alpha}{\sigma}\left[\frac{L_n\left(1-e^{-\sigma}\ \tau^{1-\sigma}\ \right)}{m+(e\tau)^{1-\sigma}}+\frac{L_s\left(e^{\sigma}\ \tau^{1-\sigma}\ -1\right)}{1+m\left(e^{-1}\tau\right)^{1-\sigma}}\right]. \tag{2.5}$$

Note that $H(0,e,\sigma,\alpha)>0$ if and only if

$$e^{\sigma-1}\ell\tau^{2\sigma-2}-\left(e^{-1}\ell+1\right)\tau^{\sigma-1}+e^{\sigma}>0. \tag{2.6}$$

Let $x=\tau^{\sigma-1},$ and $\Gamma_a(x)\equiv e^{\sigma-1}\ell x^2-\left(e^{-1}\ell+1\right)x+e^{\sigma}.$ Hence, (2.6) holds if and only if $\Gamma_a(x)>0.$ $\Gamma_a(x)$ is a parabola opening upward with $\Gamma_a(1)=-e^{-1}\ell+e^{\sigma}-1<0.$ Hence, there are two positive roots to $\Gamma_a(x)=0,$ and one is less than 1 and one is greater than 1. Since $\tau>1,$ there is only one τ satisfying $e^{\sigma}\left[\tau^{\sigma-1}\left(e^{-1}\ell\right)+\tau^{1-\sigma}\right]=e^{-1}\ell+1.$ Denote this value of τ as $\tau_a.$ So, $H(0,e,\sigma,\alpha)>0$ holds if and only if $\tau>\tau_a.$

We now show that there exists a unique positive number $\hat{m}$ defined as the solution to $\dfrac{1+m\left(e^{-1}\tau\right)^{1-\sigma}}{m+(e\tau)^{1-\sigma}}=\dfrac{1-e^{\sigma}\tau^{1-\sigma}}{1-e^{-\sigma}\tau^{1-\sigma}}\dfrac{L_s}{L_n}$ such that $H>0$ for $m<\hat{m},$ and $H<0$ for $m>\hat{m}$. Observe from (5) that $H>0$ if and only if

$$\frac{1+m\left(e^{-1}\tau\right)^{1-\sigma}}{m+(e\tau)^{1-\sigma}} > \frac{1-e^{\sigma}\tau^{1-\sigma}}{1-e^{-\sigma}\tau^{1-\sigma}}\frac{L_s}{L_n} \equiv G, \qquad (2.7)$$

the right-hand side of which is positive because $e^{\sigma}\tau^{1-\sigma} < e^{-\sigma}\tau^{1-\sigma} < 1$. The second inequality holds because $\tau > \tau_a$ implies that $\tau > e^{-\frac{\sigma}{\sigma-1}}$, because $H(0,e,\sigma,\alpha) < 0$ when $\tau = e^{-\frac{\sigma}{\sigma-1}}$. The left-hand side of (7) strictly decreases in m from $e^{\sigma-1}\tau^{\sigma-1}$ at $m=0$ to $e^{\sigma-1}\tau^{1-\sigma}$ when $m \to \infty$. If $G \geq e^{\sigma-1}\tau^{\sigma-1}$, then $H \leq 0$ for all $m \geq 0$, which contradicts that $H > 0 \; at \; m=0$. So, $\tau > \tau_a$ guarantees that $G < e^{\sigma-1}\tau^{\sigma-1}$. If $G \leq e^{\sigma-1}\tau^{1-\sigma}$, then $H > 0$ for all $m \geq 0$, and we also have $F > 0$ for all $m \geq 0$. In this case, equilibrium is such that $M_s = 0$ so that all firms are located in the North. To rule out this scenario, we must also impose that $G > e^{\sigma-1}\tau^{1-\sigma}$, which is equivalent to

$$\tau^{2\sigma-2} - \left(e^{\sigma} + \ell e^{\sigma-1}\right)\tau^{\sigma-1} + \ell e^{-1} > 0.$$

Again, let $x = \tau^{\sigma-1}$ and $\Gamma_b(x) \equiv x^2 - \left(e^{\sigma} + \ell e^{\sigma-1}\right)x + \ell e^{-1}$. $\Gamma_b(x)$ is a parabola opening upward with its minimum at $(e^{\sigma} + \ell e^{\sigma-1})/2$.

$$\Gamma_b\left(\frac{e^{\sigma} + \ell e^{\sigma-1}}{2}\right) = \ell e^{-1} - \frac{\left(e^{\sigma} + \ell e^{\sigma-1}\right)^2}{4}$$

Note that $\Gamma_b(1) = \left(1-e^{\sigma}\right)\left(1+\ell e^{-1}\right) > 0$. We distinguish the following cases. If $\frac{e^{\sigma} + \ell e^{\sigma-1}}{2} \leq 1$, then $\tau \geq 1$ is at the strictly increasing portion of the parabola. Since $\Gamma_b(1) > 0, \Gamma_b(x) > 0$ for all $\tau > 1$. So, we don't need any extra restriction in this case. If $\frac{e^{\sigma} + \ell e^{\sigma-1}}{2} > 1$ and $\Gamma_b\left(\frac{e^{\sigma} + \ell e^{\sigma-1}}{2}\right) > 0$, then

$\Gamma_b(x) > 0$ for all $\tau > 1$. But if $\frac{e^{\sigma} + \ell e^{\sigma-1}}{2} > 1$ and $\Gamma_b\left(\frac{e^{\sigma} + \ell e^{\sigma-1}}{2}\right) \leq 0$, then $\Gamma_b(x) > 0$ for all $\tau > 2^{-\frac{1}{\sigma-1}}\left[e^{\sigma} + \ell e^{\sigma-1} + \sqrt{\left(e^{\sigma} + \ell e^{\sigma-1}\right)^2 - 4\ell e^{-1}}\right]^{\frac{1}{\sigma-1}}$. In sum, we can define

$$\tau_b \equiv \begin{cases} \max\left\{1, 2^{\frac{1}{1-\sigma}}\left[e^{\sigma}\left(1+\ell e^{-1}\right) + \sqrt{e^{2\sigma}\left(1+\ell e^{-1}\right)^2 - 4\ell e^{-1}}\right]^{\frac{1}{\sigma-1}}\right\} & \text{if } 4\ell e^{-1} \leq e^{2\sigma}\left(1+\ell e^{-1}\right)^2 \\ 1 & \text{if } 4\ell e^{-1} \leq e^{2\sigma}\left(1+\ell e^{-1}\right)^2 \end{cases}$$

Thus, $G > e^{\sigma-1}\tau^{1-\sigma}$ for all $\tau > \tau_b$. Let $\hat{\tau}_i = \max\{\tau_{ai}, \tau_{bi}\}$, where τ_{ai} and τ_{bi} are the values of τ_a and τ_b when $\sigma = \sigma_i$. Also let $\hat{\tau} = \max\{\hat{\tau}_1, \hat{\tau}_2\}$. Hence, when $\tau > \hat{\tau}$,

$$e^{\sigma_i - 1}\tau^{\sigma_i - 1} > G_i > e^{\sigma_i - 1}\tau^{1-\sigma_i}, \tag{2.8}$$

and a finite $\hat{m}_i > 0$ is the unique solution to $\frac{1 + m_i\left(e^{-1}\tau\right)^{1-\sigma_i}}{m_i + (e\tau)^{1-\sigma_i}} = G_i$. Then, $H_i \equiv H(m, e, \sigma_i, \alpha_i) > 0$ for $m > \hat{m}_i$.

Now, denote $\hat{m}_{\max} = \max\{\hat{m}_1, \hat{m}_2\}$ and $\hat{m}_{\min} = \min\{\hat{m}_1, \hat{m}_2\}$. Since $F = H_1 + H_2$, $F > 0$ for $m \in [0, \hat{m}_{\min}]$, and $F > 0$ for $m \in [\hat{m}_{\max}, \infty)$. By continuity, any equilibrium m^* such that $F(m^*, \ldots) = 0$ must be in $(\hat{m}_{\min}, \hat{m}_{\max})$. Moreover, if m^* is unique, then we have $\partial F / \partial m < 0$ at m^*. Since $\partial F / \partial e > 0$ (because $\partial H / \partial e > 0$) and $\partial F / \partial m < 0$ we have, at m^*,

$$\frac{dm}{de} = -\frac{\frac{\partial F}{\partial e}}{\frac{\partial F}{\partial m}} > 0.$$

The rest of the proof proves the uniqueness of m^*, and for this purpose, we use the Descartes' rule of signs to show that there is exactly one positive root of $F=0$. Observe that

$$F=\frac{Am^3+Bm^2+Cm+D}{\left[m+(e\tau)^{1-\sigma_1}\right]\left[1+m\left(e^{-1}\tau\right)^{1-\sigma_1}\right]\left[m+(e\tau)^{1-\sigma_2}\right]\left[1+m\left(e^{-1}\tau\right)^{1-\sigma_2}\right]},$$

where

$$\begin{aligned}
&A=K_1\left(e^{-1}\tau\right)^{1-\sigma_2}+K_2\left(e^{-1}\tau\right)^{1-\sigma_1}\\
&B=P_1\left(e^{-1}\tau\right)^{1-\sigma_2}+P_2\left(e^{-1}\tau\right)^{1-\sigma_1}+K_1\left(1+\tau^{2-2\sigma_2}\right)+K_2\left(1+\tau^{2-2\sigma_1}\right)\\
&C=P_1\left(1+\tau^{2-2\sigma_2}\right)+P_2\left(1+\tau^{2-2\sigma_1}\right)+K_1(e\tau)^{1-\sigma_2}+K_2(e\tau)^{1-\sigma 1}\\
&D=P_1(e\tau)^{1-\sigma_2}+P_2(e\tau)^{1-\sigma_1},
\end{aligned}$$

where

$$\begin{aligned}
K_i&=\frac{\alpha_i}{\sigma_i}\left[L_n\left(1-e^{-\sigma_i}\tau^{1-\sigma_i}\right)\left(e^{-1}\tau\right)^{1-\sigma_i}+L_s\left(e^{\sigma_i}\tau^{1-\sigma_i}-1\right)\right]\\
P_i&=\frac{\alpha_i}{\sigma_i}\left[L_n\left(1-e^{-\sigma_i}\tau^{1-\sigma_i}\right)\left(e^{-1}\tau\right)^{1-\sigma_i}+(e\tau)^{1-\sigma_i}L_s\left(e^{\sigma_i}\tau^{1-\sigma_i}-1\right)\right]
\end{aligned}$$

Since $F=0$ if and only if and only if $Am^3+Bm^2+Cm+D=0$, it suffices to show that there is exactly one positive root to this polynomial. By (8), we know that

$$\begin{aligned}
K_i&=\frac{\alpha_iL_n\left(1-e^{-\sigma_i}\tau^{1-\sigma_i}\right)}{\sigma_i}\left[e^{\sigma_i-1}\tau^{1-\sigma_i}-G\right]<0,\\
P_i&=\frac{\alpha_iL_n\left(1-e^{-\sigma_i}\tau^{1-\sigma_i}\right)(e\tau)^{1-\sigma_i}}{\sigma_i}\left[e^{\sigma_i-1}\tau^{1-\sigma_i}-G\right]>0.
\end{aligned}$$

Hence, $A<0$, $D>0$, and

$$B=\left[\left(e^{-1}\tau\right)^{1-\sigma_2}+\left(1+\tau^{2-2\sigma_2}\right)\frac{e^{\sigma_1-1}\tau^{1-\sigma_1}-\dfrac{1-e^{\sigma_1}\tau^{1-\sigma_1}}{1-e^{-\sigma_1}\tau^{1-\sigma_1}}\dfrac{L_s}{L_n}}{e^{\sigma_1-1}\tau^{\sigma_1-1}-\dfrac{1-e^{\sigma_1}\tau^{1-\sigma_1}}{1-e^{-\sigma_1}\tau^{1-\sigma_1}}\dfrac{L_s}{L_n}}\left(e\tau\right)^{\sigma_1-1}\right]P_1$$

$$+\left[\left(e^{-1}\tau\right)^{1-\sigma_1}+\left(1+\tau^{2-2\sigma_1}\right)\frac{e^{\sigma_2-1}\tau^{1-\sigma_2}-\dfrac{1-e^{\sigma_2}\tau^{1-\sigma_2}}{1-e^{-\sigma_2}\tau^{1-\sigma_2}}\dfrac{L_s}{L_n}}{e^{\sigma_2-1}\tau^{\sigma_2-1}-\dfrac{1-e^{\sigma_2}\tau^{1-\sigma_2}}{1-e^{-\sigma_2}\tau^{1-\sigma_2}}\dfrac{L_s}{L_n}}\left(e\tau\right)^{\sigma_2-1}\right]P_2$$

$$C=\left[1+\tau^{2-2\sigma_2}+\left(e\tau\right)^{1-\sigma_2}\frac{e^{\sigma_1-1}\tau^{1-\sigma_1}-\dfrac{1-e^{\sigma_1}\tau^{1-\sigma_1}}{1-e^{-\sigma_1}\tau^{1-\sigma_1}}\dfrac{L_s}{L_n}}{e^{\sigma_1-1}\tau^{\sigma_1-1}-\dfrac{1-e^{\sigma_1}\tau^{1-\sigma_1}}{1-e^{-\sigma_1}\tau^{1-\sigma_1}}\dfrac{L_s}{L_n}}\left(e\tau\right)^{\sigma_1-1}\right]P_1$$

$$+\left[1+\tau^{2-2\sigma_1}+\left(e\tau\right)^{1-\sigma_1}\frac{e^{\sigma_2-1}\tau^{1-\sigma_2}-\dfrac{1-e^{\sigma_2}\tau^{1-\sigma_2}}{1-e^{-\sigma_2}\tau^{1-\sigma_2}}\dfrac{L_s}{L_n}}{e^{\sigma_2-1}\tau^{\sigma_2-1}-\dfrac{1-e^{\sigma_2}\tau^{1-\sigma_2}}{1-e^{-\sigma_2}\tau^{1-\sigma_2}}\dfrac{L_s}{L_n}}\left(e\tau\right)^{\sigma_2-1}\right]P_2$$

So, because $1+\tau^{2-2\sigma_i}>1>\left(e^{-1}\tau\right)^{1-\sigma_i}$ and $1+\tau^{2-2\sigma_i}>1>\left(e\tau\right)^{1-\sigma_i}$, $B<C.$ To apply the rule of signs, distinguish three cases, $C>B>0, 0>C>B,$ and $C>0>B.$ Combined, with the facts that $A<0$ and $D>0,$ there is exactly one positive root in each case. Hence, m^* is unique.

Proof of Proposition 2

Observe that $\dfrac{d}{de}\left(\dfrac{X_{s2}}{X_{s1}}\right)>0$ if and only if

$$\frac{d}{de}\left(\frac{P_{n1}^{1-\sigma_1}}{P_{n2}^{1-\sigma_2}}e^{\sigma_1-\sigma_2}\right)=\frac{d}{de}\left(\frac{\lambda_1\left(\mu_1 c\right)^{1-\sigma_1}\left[m+\left(e\tau\right)^{1-\sigma_1}\right]}{\lambda_2\left(\mu_2 c\right)^{1-\sigma_2}\left[m+\left(e\tau\right)^{1-\sigma_2}\right]}e^{\sigma_1-\sigma_2}\right)>0,$$

which is positive if and only if

$$\frac{d}{de}\left(\frac{m+\left(e\tau\right)^{1-\sigma_1}}{m+\left(e\tau\right)^{1-\sigma_2}}e^{\sigma_1-\sigma_2}\right)>0.$$

In the short run when m is kept fixed, we have

$$\frac{\partial}{\partial e}\left(\frac{m+(e\tau)^{1-\sigma_1}}{m+(e\tau)^{1-\sigma 2}}e^{\sigma_1-\sigma_2}\right)$$
$$=\frac{e^{-1}\tau^{\sigma_2-\sigma_1}\left[m^2(e\tau)^{\sigma_1+\sigma_2}(\sigma_1-\sigma_2)+m(e\tau)^{\sigma_1+1}(\sigma_1-1)-m(e\tau)^{\sigma_2+1}(\sigma_2-1)\right]}{\left[e\tau+m(e\tau)^{\sigma_2}\right]^2}.$$

It is easy to verify that $\hat{\tau}>e^{-1}$ by checking (6) and so $e\tau>1$. Hence, the above derivative is positive. Next, in the long run when m is not fixed, we must examine

$$\frac{\partial}{\partial m}\left(\frac{m+(e\tau)^{1-\sigma_1}}{m+(e\tau)^{1-\sigma_2}}e^{\sigma_1-\sigma_2}\right)=\frac{e\tau^{1-\sigma_1+\sigma_2}\left[(e\tau)^{\sigma_1}-(e\tau)^{\sigma_2}\right]}{\left[e\tau+m(e\tau)^{\sigma_2}\right]^2},$$

Which is positive since $e\tau>1$. From Proposition 1, $\frac{dm}{de}>0,$ and hence at m^*,

$$\frac{d}{de}\left(\frac{m+(e\tau)^{1-\sigma_1}}{m+(e\tau)^{1-\sigma 2}}e^{\sigma_1-\sigma_2}\right)=\frac{\partial}{\partial e}\left(\frac{m+(e\tau)^{1-\sigma_1}}{m+(e\tau)^{1-\sigma_2}}e^{\sigma_1-\sigma_2}\right)+\frac{\partial}{\partial m}\left(\frac{m+(e\tau)^{1-\sigma_1}}{m+(e\tau)^{1-\sigma_2}}e^{\sigma_1-\sigma_2}\right)\frac{dm}{de}>0.$$

Hence, $\frac{d}{de}\left(\frac{X_{s2}}{X_{s1}}\right)>0.$ The proof for $\frac{d}{de}\left(\frac{X_{n2}}{X_{n1}}\right)<0$ is similar.

For Point 2, first observe from (2.2) and (2.3), $\frac{X_{n2}}{X_{n1}}>\frac{X_{s2}}{X_{s1}}$ if and only if

$$\frac{P_{s1}^{1-\sigma_1}}{P_{s2}^{1-\sigma_2}}\left(e^{-1}\right)^{\sigma_1-\sigma_2}>\frac{P_{n1}^{1-\sigma_1}}{P_{n2}^{1-\sigma_2}}e^{\sigma_1-\sigma_2},$$

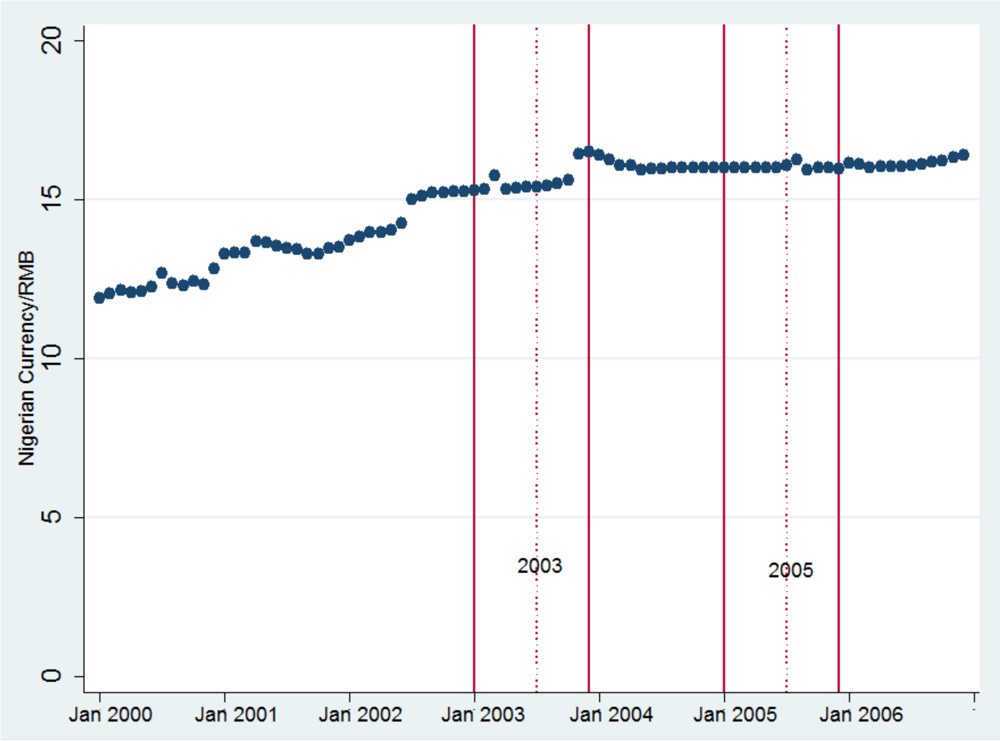
If and only if

$$\frac{1+m\left(e^{-1}\tau\right)^{1-\sigma_1}}{1+m\left(e^{-1}\tau\right)^{1-\sigma_2}}\left(e^{-1}\right)^{\sigma_1-\sigma_2}>\frac{m+(e\tau)^{1-\sigma_1}}{m+(e\tau)^{1-\sigma_2}}\left(e^{-1}\right)^{\sigma_1-\sigma_2}. \qquad (2.9)$$

We can first look at a special case where $e=1$. The only difference between two countries in this case is the scenario when $L_n\neq L_s$. When

$L_n > L_s$, $m = \frac{M_n}{M_s} < 1$. With $e = 1$ and $m < 1$, it is easy to verify that (2.9) holds. This proves Point 2(a), let $L_n = L_s$. When $e = 1, m = 1$ and (2.9) holds in equality, i.e., $\frac{X_{s2}}{X_{s1}} = \frac{X_{n2}}{X_{n1}}$. By Point 1, when $e < 1, \frac{X_{s2}}{X_{s1}} < \frac{X_{n2}}{X_{n1}}$ holds.

Appendix Figure 2.1



Appendix 3 for Chapter Three

Appendix Table 3.1: More Robustness Checks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable	Δ Output Share of SOEs						
Specification	Alternative SOE definition	Two-periods	Employment -share	Sales 5m+	Excl. foreign firms	Domestic sale	1997-2001 tariffs
ΔTariff	-0.277*	-0.294**	-0.187**	-0.304***	-0.254**	-0.265**	-0.307***
	(0.147)	(0.149)	(0.085)	(0.107)	(0.123)	(0.108)	(0.112)
Province-Year Fixed Effects	X	X	X	X	X	X	X
Δ Ln(GDP per capita)	X	X	X	X	X	X	X
Δ Gov Consumption/GDP	X	X	X	X	X	X	X
Δ Ln(Dairy per Capita)*Post02	X	X	X	X	X	X	X
Δ Ln(Vegetable per Capita)*Post02	X	X	X	X	X	X	X
Δ Ln(Telephone)*Post02	X	X	X	X	X	X	X
Δ University*Post02	X	X	X	X	X	X	X
Δ Coastal*Post02	X	X	X	X	X	X	X
Δ Northern*Post02	X	X	X	X	X	X	X
Δ Mountain*Post02	X	X	X	X	X	X	X
Number of Observations	1697	252	1697	1697	1697	1697	1697
R-squared	0.2707	0.3791	0.2387	0.2292	0.2220	0.1928	0.2289

Notes: Standard errors, clustered at the city level, are reported in the parenthesis. *** p<0.01, ** p<0.05, * p<0.1.