Labor, Trade and Finance

Essays in Applied Economics

Mengyi Cao

Academic dissertation for the Degree of Doctor of Philosophy in Economics at Stockholm University to be publicly defended on Friday 15 December 2017 at 15.00 in De Geersalen Geovetenskapens hus, Svante Arrhenius väg 14.

Abstract

Essay I: Credit Constraint and College Attendance.

This paper shows that housing wealth alleviate credit constraints for potential college attendees by enabling home owners to extract equity from their property and invest it in the education. Using a large US individual-level survey dataset over the 1996-2011 period, I find that one standard deviation increases of housing prices translate into approximately 72,000 more students enrolled in college each year. My results stay significant when I use proxies for aggregate housing demand shocks and for the topological elasticity of housing supply to generate variation in home equity that is assumed to be orthogonal to decision of going to college.

Essay II: Income Inequality and Trade.

Does trade with unskilled labor-abundant countries reduce the relative wages of U.S. unskilled labor and consequently cause increased income inequality across industries and regions? Empirical studies in the 1990s found only a modest effect. In this paper, I re-consider the question by using the income inequality measures constructed from Current Population Survey (CPS) data and analyzing the effect of rising Chinese import competition between 1993 and 2007 on US local labor markets. I find that areas which are more exposed to China imports competition have larger changes in income inequality. In my main specification, a $1,000 exogenous decadal rise in a MSA's import exposure per worker leads to a 1.5% increase in the logistic Gini. This re-distributive effect is more profound among non-college educated workers in manufacturing sectors.

Essay III: Employee as Creditor: Evidence from Defined Pension Plans.

In this paper, I show the role of pension plans in shaping the firms' labor market decision. By employing the loan covenants violation and consequently transferring of control rights to creditors, I examine the strategic use of pension underfunding by firms and the resultant wage cuts. I also find that the wage concession is less severe for firms from industry with bigger bargaining power. This study sheds light on how firms strategically renegotiate labor contracts to extract concessions from labor. The evidence suggests that credit contracts between debt-holders and shareholders have spillover effects on non-financial stakeholders.

Keywords: *Human Capital, Education, Home Equity, International Trade, Pension Plan.*

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Department of Economics

Stockholm University, 106 91 Stockholm

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Mengyi Cao

Stockholm
University

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List of Papers

The following essays are included in this thesis.

Essay I: Credit Constraint and College Attendance

Essay II: Income Inequality and Trade

Essay III: Employee as Creditor: Evidence from Defined Pension Plans

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Stockholm, November, 2017
Mengyi Cao

Introduction

This Ph.D. thesis consists of three essays in empirical studies on economics. Short summaries of the essays are provided below.

Essay I: Credit Constraint and College Attendance

This paper shows that housing wealth alleviate credit constraints for potential college attendees by enabling home owners to extract equity from their property and invest it in the education. Using a large US individual-level survey dataset over the 1996-2011 period, I find that one standard deviation increases of housing prices translate into approximately 72,000 more students enrolled in college each year. My results stay significant when I use proxies for aggregate housing demand shocks and for the topological elasticity of housing supply to generate variation in home equity that is assumed to be orthogonal to decision of going to college.

The results of the paper have particular relevance to current policy as credit markets have tightened and housing prices have declined in many areas of the country. Considering the reduction in family resources caused by these problems in the housing market, it is likely that many families will face increasing constraints in their ability to finance college in the near future. This work provides new evidence on that college attendance is sensitive to these fluctuations.

Essay II: Income Inequality and Trade

Does trade with unskilled labor-abundant countries reduce the relative wages of U.S. unskilled labor and consequently cause increased income inequality across industries and regions? Empirical studies in the 1990s found only a modest effect. In this paper, I re-consider the question by using the income inequality measures constructed from Current Population Survey (CPS) data and analyzing the effect of rising Chinese import competition between 1993 and 2007 on US local labor markets. I find that areas which are more exposed to China imports competition have larger changes in income inequality. In my main specification, a $1,000 exogenous decadal rise in a MSA's import exposure per worker leads to a 1.5% increase in the logistic Gini. This re-distributive effect is more profound among non-college educated workers in manufacturing sectors.

Essay III: Employee as Creditor: Evidence from Defined Pension Plans

In this paper, I show the role of pension plans in shaping the firms' labor market decision. By employing the loan covenants violation and consequently transferring of control rights to creditors, I examine the strategic use of pension underfunding by firms and the resultant wage cuts. I also find that the wage concession is less severe for firms from industry with bigger bargaining power. This study sheds light on how firms strategically renegotiate labor contracts to extract concessions from labor. The evidence suggests that credit contracts between debt-holders and shareholders have spillover effects on non-financial stakeholders.

Sammanfattning på svenska

Denna doktorsavhandling består av tre uppsatser som behandlar empiriska studier i nationalekonomi. Korta sammanfattningar av uppsatserna återfinns nedan.

Uppsats 1: Credit Constraint and College Attendance (Kreditbegränsning och universitetsstudier)

Denna uppsats visar att förmögenhet i form av bostäder mildrar kreditbegränsningarna för potentiella universitetsstudenter genom att möjliggöra för bostadsägare att utnyttja nettovärde av sin egendom och investera detta i utbildning. Genom att använda ett stort dataset från en amerikansk undersökning påindividnivåunder perioden 1996-2001, finner jag att en ökning i bostadspriserna motsvarande en standardavvikelse motsvarar ca 72000 ytterligare studenter som skrivs in vid universitetet varje år. Mina resultat förblir signifikanta när jag använder s k *proxies* för aggregerade störningar vad gäller efterfrågan påbostäder och den topologiska elasticiteten när det gäller utbudet påbostäder för att generera en variation i bostadstillgångar som antas vara ortogonal med beslutet att gåpåuniversitetet.

Resultaten i uppsatsen är speciellt viktiga för den nuvarande politiken dåkreditmarknaderna har stramats åt och bostadspriserna har sjunkit i många områden i landet. Med tanke påden minskning av familjeresurserna som skett till följd av dessa problem påbostadsmarknaden, kommer sannolikt många familjer att ståinför ökande begränsningar när det gäller deras förmåga att finansierar universitetsstudier den närmaste framtiden. Denna uppsats ger nya bevis för att deltagandet i universitetsstudier är känsligt för dessa fluktuationer.

Uppsats 2: Income Inequality and Trade (Inkomstomjämlikhet och handel)

Minskar handeln med länder som har gott om okvalificerad arbetskraft relativlönerna för okvalificerad arbetskraft i USA och skapar följaktligen en ökad inkomstojämlikhet mellan branscher och regioner? Empiriska studier som utfördes på1990-talet fann endast en liten effekt. I den här uppsatsen tar jag pånytt upp frågan genom att använda mått påinkomstomjämlikhet som skapas från Current Population Survey Data (CPS) och analyserar effekten av en ökande konkurrens från kinesisk import mellan 1993 och 2007 pålokala amerikanska arbetsmarknader. Jag finner att de områden som är mest utsatta för importkonkurrens från Kina har större förändringar i inkomstomjämlikheten. I min huvudsakliga specifikation leder en exogen ökning uppgående till USD 1000 per decennium i en MSAs (Metropolitan Statistical Area) utsatthet för import per anställd till en 1,5% ökning i logistisk Gini. Denna omfördelningseffekt är djupare bland anställda i tillverkningssektorerna som saknar en universitetsutbildning.

Uppsats 3: Employee as Creditor: Evidence from Defined Pension Plans (Anställd som borgenär: Bevis från definierade pensionsplaner)

I denna uppsats visar jag pensionsplanernas roll när det gäller att utforma företagens arbetsmarknadsbeslut. Genom att använda överträdelser av låneavtal och följaktligen överföring av kontrollrätter till borgenärer, undersöker jag den strategiska användningen av underfinansiering av pensioner av företagen och de därav följande lönesänkningarna. Jag finner även att löneminskningarna är mindre allvarliga i branscher med större förhandlingsutrymme. Denna studie belyser hur företag strategiskt omförhandlar anställningskontrakt för att fåmedgivande till försämringar i ersättningen för arbete. Bevisen tyder påatt kreditavtal mellan obligationsinnehavare och aktieinnehavare har överspillningseffekter påickefinansiella intressenter som satsat pengar.

1. Credit Constraint and College Attendance

This paper shows that housing wealth alleviate credit constraints for potential college attendees by enabling home owners to extract equity from their property and invest it in the education. Using a large US individual-level survey dataset over the 1996-2011 period, I find that one standard deviation increases of housing prices translate into approximately 72,000 more students enrolled in college each year. My results stay significant when I use proxies for aggregate housing demand shocks and for the topological elasticity of housing supply to generate variation in home equity that is assumed to be orthogonal to decision of going to college.

1.1 Introduction

Human capital is crucial for technological change which is the engine of economic growth, and postsecondary enrollment is the key to foster human capital accumulation. There is increasing interest on understanding the impact of credit constraints on the accumulation of human capital. Credit constraint resulting from imperfect capital markets could affect an individual's investment in human capital. Recent studies appear to suggest that borrowing constraints may be harmful for investment decisions in young children (Lochner and Naranjo (2012) and also a more recent study by Solis (2013)).

The prior studies on analyzing the effect of credit constraints on investment in education are challenging for at least two reasons. First, lack of wealth information in the data sets gives an incomplete picture as if families use both income and wealth to pay for college. Second, even if wealth measures are available, establishing the causality between wealth and college enrollment is still difficult because wealthier families more likely to send their children to college, and this tendency is likely at least partly attributed to unobserved factors such as child ability and preferences for education.

This paper complements the existing literature by analyzing if and how a particular wealth, housing wealth, affects college attendance. In the presence of imperfect capital markets (credit constraints), families can use home equity as collateral to finance the education investments. Through this collateral channel, shocks to the value of real estate can have a large impact on college attendance decisions for those students who are credit constrained. This paper makes three main contributions: First, instead of assuming an exogenous family resource, I use local variations in house prices as shocks to collateral value of house owners. More precisely, I instrument the home equity with the interaction of local housing supply elasticity and house price growth. Second, it is the first formal analysis of the effect of collateral channel on educa-

tional investment decision. Previously, the collateral channel has been argued as being of first-order importance to firm investment (Chaney et al.(2011)), firm capital structure (Cvijanovic (2014)), entrepreneurship (Corradin and Popov (2015)) and innovation (Cao et al.(2014)). Third, while many existing studies have emphasized the effect of student aid, family income or wealth, my paper complements the existing literature by considering another important source of financing to household, namely the home equity borrowing channel from banks.

The most relevant studies to this paper are Dynarski (2002), Lovenheim(2011) and Lovenheim and Reynolds (2013). For instance, Lovenheim(2011) uses the variation in housing prices during 2000s as an instrument for housing wealth to identify the effect of household wealth on college enrollment and find a positively significant effect of housing wealth. There are several distinctions between my paper and theirs: firstly, as also mentioned in Lovenheim(2011), their methodology is not able to parse out the wealth effect of house value from the credit effect. Because their findings could be likely confounded by an increase in house wealth which simultaneously leads more consumption goods including college education. The empirical strategy I employ is credible because using interaction of land supply elasticity and house price growth between the current year and the year of purchase as an instrument on housing prices could potentially capture the effect of credit channel well. Secondly, their work is primarily focused on the housing boom period. For example, Lovenheim(2011) and Lovenheim and Reynolds (2013) only focus on years 2001-2005 while Dynarski (2002)'s data is limited in early 1990s. This is incomplete given the housing bubble bust and the followed economics downturn could also give rich implications to the question we look at. Thus instead of merely focusing on a relatively short period, the data I use in the paper has a longer time span from 1996-2011.

This paper is related to a growing research area, mainly on corpo-

rate and household finance, that analyses the effect collateral channel on corporate investment (Chaney et al.(2011)) and debt capacity (Cvijanovic (2014)), bank loans (Cunant et al.(2014)), household leverage level (Mian and Sufi (2011)), households' risky asset portfolio (Chetty and Szeidl (2010)). It is also related to those analyzing the effect of collateral channel on individual investment, consumption and re-financing behavior (Case, Quigley, and Shiller (2005), Bostic, Gabriel, and Painter (2009), Lehnert (2004), Hurst and Stafford (2004), Campbell and Cocco (2007), Greenspan and Kennedy (2007)). My paper is among the first ones that look at how housing wealth impacts investment on education by utilizing the identification strategy which clearly separates the credit effect.

1.2 Background

There are two non-mutually exclusive theoretic arguments on how a family's variation in its housing wealth to influence its children's education decisions. One perspective treats college education as an investment good as in the model of Becker (1962). In Becker and Tomes (1979,1986), they develop a model which is based on the assumption that parents are utility-maximum driven and investing in the future success of their children. This line of research argues that parents invest in college only if the net return exceeds or equals to the market rate of return from forgone earnings getting the college education and the direct cost of college. Family resources, and housing wealth in particular, have the possibility of affecting college enrollment through the credit channel: students who are credit-constrained can borrow from banks via their parents who can use their house as collateral to obtain loans at a lower interest rate, at the same time, banks could begin lending money to those families which previously excluded from the credit market.

An alternative perspective sees college enrollment as a consumption

good (Lazear (1979)). In this case, college enrollment increases due to wealth effect. Lovenheim (2011) points out that as home prices grow and homeowners become richer, they would consume more good which might include college education. Thus housing wealth variation can impact college enrollment decision in the absent of credit constraints through wealth effect channel instead of credit channel. In short, the positive relationship between housing variation and college attendance might simply be due to the increased household wealth.

The challenge of previous studies is to distinguish and test these two hypotheses. The factors that affect credit constraints of a family are usually associated with its wealth level. For instance, Lovenheim (2011) employs the short-run changes in home prices as an instrument for the variation in housing wealth. However, the change of house prices naturally involve wealth effects in the estimates. Areas which undertake a faster house price growth could have also been experiencing a good economic growth. Thus a higher college attendance rate in that area might simply be reflecting the increase of family income or a better off local economic condition.

The central part of the paper deals with the identification of the credit effect through the collateral channel of housing wealth. As mentioned above, there are non-trivial endogeneity concerns related to the impact of home equity on the transition into college attendance. For instance, they may disproportionately reside in areas with booming local economies where the propensity to go to college is higher. In fact, it is highly likely that the areas which experienced the largest house price booms during the early-to-mid 2000s were also intrinsically areas with higher college attendance rate (such as large coastal cities, more in Table 4). Alternatively, an outward shift in the supply of credit which accelerated house price growth (see Mian and Sufi 2011) may have also relaxed constraints on student loans, leading to higher levels of college attendance.

I implement a number of empirical strategies to differentiate the exogenous shocks to housing wealth from local economic effect (such as demand booms) which may be correlated with changes in housing prices and in college attendance. First, following Chetty and Szeidl (2010) and Mian and Sufi (2011)[1], I instrument the home equity using the change in average national house prices between the year when the house was purchased and the current year, times the local housing supply elasticity. The idea behind this approach is that an increase in the economy-wide demand for housing will increase house prices, and this effect should be stronger in MSAs with less elastic housing supply where the adjustment in response to aggregate demand shocks takes place on the price margin. Intuitively, for a location with a very high elasticity of land supply, an increase in demand will be fully translated into increased quantity of new construction instead of higher real estate prices. For a location with inelastic land supply, because the new construction is constrained by geography, an increase in demand will be translated into higher housing prices. Second, I conduct a couple of robustness checks to account for the underlying mechanism of the credit channel. I repeat the analysis on sub-sample of individual comes from high and low level of family income. By doing so, I test the hypothesis that if the effect is stronger from families that are more financially constrained. And also I repeat the analysis on sub-sample in different years to examine whether the effect of housing wealth varies alongside with business cycle. Third, I compare home owners to renters, hypothesizing that the same increase

[1]This IV methodology has many variations since it is first introduced in Himmelberg, Mayer, and Sinai (2005). For instance, Mian and Sufi (2011) instrument house price growth with the interaction of zip code level subprime share and MSA level housing supply inelasticity. Similar to Mian and Sufi (2011) is Chaney et al.(2011) which instrument housing price with the interaction of national interest rates and inelasticity. In another version of the identification strategy, Chetty and Szeidl (2010) and also Corradin and Popov (2015), they use the change in average *US-wide* house prices times the local elasticity of housing supply as an instrument for home equity.

in house prices should affect college attendance propensity relatively more for home owners as renters lack entirely the collateral channel.

1.3 Empirical Strategy

I start with different specifications of a simple reduced form regression to analyze the impact of housing wealth variation on college enrollment. Specifically, I run the following regression:

$$Enroll_{it} = \beta_0 + \beta_1 Own_{it} + \beta_2 HouseValue_{it} + \gamma X_i + l_i + t_i + \varepsilon_i, \quad (1.1)$$

where $Enroll_{it}$ is an indicator of whether or not an individual i has enrolled in a 2-year or 4-year college, Own_{it} is a dummy variable equals to one if the household owns its home in time t, $HouseValue_{it}$ refers to the home equity of household i, X_i include a vector of individual and household characteristics such as household's current income and net wealth. Ignoring the effect of wealth may bias our results upward because individuals who come from richer family more likely afford to attend college. Finally, I include year fixed effects t_i and geographic area fixed effects l_i.

While the econometric specification allows us to study the relationship between current home equity and college attendance decisions, it may still suffer from potential omitted variable bias. Individuals who ultimately choose to attend college may disproportionately reside in areas with local economies or tradition where the propensity to attend college is naturally higher. The nature of local economies or the cultural norms regarding the propensity of college attendance is difficult to observe and control. House value or local house price are affected by those hidden factors. For instance, better investment opportunities. If one sees eduction as an investment good, it is very likely families in the city which experiences a large increase of housing price would also make more investment which includes education investment. It is also intuitively

reflected at the first glance of the data: those areas which experienced the largest house price booms during the early-to-mid 2000s happen to coincide with the same areas with higher college enrollment rate (e.g. California, New York, Connecticut and Massachusetts). Therefore, even if one finds a positive association between home equity or local land price and college enrollment, it may be largely driven by those omitted variables. In all, it is crucial to identify the causality effect of home borrowing channel from the total effect of housing wealth variation on college enrollment decision.

In order to address the endogeneity of housing wealth, I instrument the home equity. Following Chetty and Szeidl (2010) and Mian and Sufi (2011), I do so by instrumenting the home equity using the change in average national house prices between the year when the house was purchased and the current year, times the local housing supply elasticity. By introducing the measure of land constraint ratio, one can mitigate the endogenity concern to some extend. First, the geographic constraint can capture the land price variation pretty well. For a location with a very high elasticity of land supply, an increase in demand will be fully translated into increased quantity of new construction instead of higher real estate prices. For a location with inelastic land supply, because the new construction is constrained by geography, an increase in demand will be translated into higher housing prices. Second, since geographic constraint is an exogenous determinant of the supply side of housing market, it is largely unaffected by any other demand side drivers of house prices. I also present a geographic motivation for this argument in Figure 1. It shows the evolution of average state residential prices from 1990 to 2011 for States with high vs low housing supply elasticity. Low elasticity states experienced a much larger increase in residential real estate prices than the high elasticity states, particularly in the early of 2000's: from 2000 to 2006, residential prices have increased by 59% while they increased only by approximately 24% for states with high

elasticity.

The advantage of this source of variation is that it avoids the potential for omitted variable bias due to local economic shocks because the variation is driven purely by national demand shocks. I calculate the elasticity both at the state level and at MSA level.[1] MSA level data is preferred while it limits my data sample because the SIPP stopped reporting the MSA code starting from 2004-2006 wave. Therefore I apply the state-level elasticity in most of the cases. I thus estimate, for State l, at date t, the following predicting house value $HouseValue_t^l$:

$$HouseValue_t^l = \delta^l + \zeta^t + \gamma Elasticity^l \times HousePriceGrowth_t + \mu_t^l, \tag{1.2}$$

where $Elasticity^l$ is the land supply elasticity at state level, $HousePriceGrowth_t$ is the national-wide house price change between the current house price and price when the property was purchased, δ^l state fixed effects and ζ^t captures the time fixed effect. This regression gives the first stage of the 2SLS model.

1.4 Data and Sample

The main source of my individual-household dataset is the Survey of Income and Program Participation (SIPP) of the U.S. Census Bureau from 1996 to 2011. Each SIPP panel tracks 20,000 to 30,000 households over a period of 2-3 years and contains information on income, assets and demographics. Four consecutive groups of households or panels were interviewed during the years 1996-2000, 2001-2004, 2004-2008 and 2008-2011. The SIPP crucially contains data on educational attainment and state of residence, as well as rich demographic controls

[1]The original data in Saiz (2010) is at the MSA level. I calculate the state-level elasticities by averaging the MSA-level elasticities for all MSAs in the same state, weighted by the MSA's population.

including household income and house value which I take as home value in the baseline regression. More specifically, the SIPP contains data on household's home value and home mortgage in the Asset and Liabilities Topic Module, which allows me to construct a proxy for home equity by taking the difference of the two. During the four panels, assets data were collected from the households interviewed during the years 1996-2000 (four times), 2001-2004 (three times), 2004-2008 (two times), and 2008-2011 (three times), respectively.

To construct the analysis sample of college enrollment, I take the following steps. First, because the SIPP are monthly data, I measure enrollment using information from September (fall enrollment). Second, following prior studies such as Dynarski (2003) and Turner (2011), I limit the sample to 18 to 19 year olds who already obtain high school diplomas, to capture college entry and the transition into the second year of college. Third, I merge the observations on individual's enrollment with the Assets Topic Module and keep those ones have relevant family-level assets data. The data covers the period from 1996-1999, 2001-2005 and 2009-2011. It contains 16,378 college-aged youths. Roughly 52% of the youths in this sample are enrolled full-time in the first two years of college, with roughly 34% enrolled in the first year and 18% in the second year. Another 8% are enrolled either part-time in the first-two years of college or enrolled in vocational, technical or business school beyond high school level. This enrollment rate is approximately the same as the rates calculated from the CPS or from the U.S. Census. Calculations from the October CPS show that college in the 2000s was about 58%.

We obtain annually data on average of housing prices by states from 1990-2011 using the repeated sales index constructed by the Office of Federal Housing Enterprise Oversight (OFHEO). Local housing supply elasticities are from Saiz (2010) where he calculates for 95 MSAs. Those MSAs, according to Saiz (2010), all have population over 500,000 in the 2000 Census and cover the most densely populated areas in the

Table 1: Summary Statistics I

This table reports the descriptive statistics for the sample of college attendance. The definition of variables are listed in Table A1. The sample is constructed from households in SIPP for the year 1996-2000, 2001-2004, 2004-2008 and 2008-2011.

	Mean	Median	SD	25^{th}	75^{th}	N
Demographics:						
College Attendance	0.61	1	0.49	0	1	16378
Age	18.5	19	0.5	18	19	16378
Female	0.51	1	0.5	0	1	16378
African American	0.14	0	0.34	0	0	16378
Housing and Wealth:						
Own	0.71	1	0.46	0	1	16378
Home Equity	74739.6	28000	123094.4	0	101000	16141
Income	5638.5	4392.5	5718.8	2100	7519	16378
Household Non-Housing Wealth	138406.6	19950	2361094.1	3694	99083	16141
Home Property Value	131613.7	90000	160841.7	0	190000	16141
Housing Prices:						
Elasticity	1.95	1.61	0.95	1.27	2.48	836
State Residential Housing Prices	265.3	237.4	111.6	189	306.1	836

U.S. Using data on physical and regulatory constrains (land availability and use regulations), this index provides a convenient index of the supply constraints at MSA level.

Table 1 reports summary statistics for the constructed enrollment sample. 14% of the college-aged youths are African-American, 51 % are female. The average amount of home equity is $ 75,000. Average monthly household income is around $ 5,638, and average non-housing wealth (i.e. total wealth excluding home equity) is around $138,400. The average property is worth $ 131,613.

Table 2 reports descriptive statistics on the subsample of college-aged youths who are enrolled in the college. I compare those to descriptive statistics on the subsample of youths who are not enrolled in the college. The differences between the two samples are in most cases statistically significant. For instance, those who are enrolled are more likely to be female and less likely to be African American. They are also more likely coming from families with higher income and more likely to own housing properties. More importantly, those who enrolled have more home equity, as well as a more valuable property. This facts seem to suggest that when estimating the effect of home equity on college attendance, we need to take care of the effect of family income and level of accumulated wealth on the probability of college attendance. Table 3 shows the comparison of descriptive statistics between sub-samples of States with low and high elasticity. Low (High) elasticity is the bottom (top) tercile of the elasticity. A simple t-test using standard error clustered at state level is conducted to test the equality of the means between two groups. We can observe that except the college enrollment, two samples are not statistically significant different from each other in youths' age, gender, race, family's ownership of house, labor income, home equity and household non-housing wealth. Table 4 reports the college enrollment by states ranked by elasticity of each state.

Table 2: Summary Statistics II

This table reports the descriptive statistics for the sub-samples of college enrolled youths and non-enrolled youths. The definition of variables are listed in Table A1. The sample is constructed from households in SIPP for the year 1996-2000, 2001-2004, 2004-2008 and 2008-2011. All statistics are means. The unweighted percentage of individuals who are enrolled in college when they were interviewed is 0.61.

	College Enrolled Youths	Non-College Enrolled Youths	N_1	N_2	p-value of difference
Demographics:					
Age	18.5	18.6	9948	6430	< 0.01
Female	0.55	0.45	9948	6430	< 0.01
African American	0.12	0.16	9948	6430	< 0.01
Housing and Wealth					
Own	0.78	0.6	9948	6430	< 0.01
Home Equity	76335.3	72285.7	9781	6360	0.04
Income	6405.3	4452.2	9948	6430	< 0.01
Household	155847.1	111585	9781	6360	0.24
Non-Housing Wealth					
Home Property Value	134033.6	127892.3	9781	6360	0.02

1.5 Empirical Results

In this section, I present different sets of empirical results. First, I present the evidence from probit regression which shows the association between housing wealth and college enrollment, together with a set of robustness checks. Second, I present the instrumental variable regression which aims to establish causality between housing wealth and college attendance.

1.5.1 The Association between Housing Wealth and College Attendance

1.5.1.1 Main Result

Table 5 reports the results from the baseline probit regression model. I firstly analyze the entire sample with all households (Column (1) to (3)) and then repeat the tests on sub-sample of home owners (Column (4) and (5)). I control for the demographic characteristics of the individual: income, gender, race alongside with home equity. To more precisely isolate the effect of housing wealth from the rest of the household's financials, I report specification without (column (1), (3) and (5)) and with (column (2) and (4)) controlling non-housing wealth. For the completeness purpose of the analysis, in column (3), I add the regression using the subsample before 2003 which contains MSA geographic information and cluster the standard errors by MSA level. The number of observations is significantly reduced due to fact that SIPP stops reporting MSA level information after 2003.

The results suggest that college enrollment and housing value is positively associated. The coefficients are persistent across different specifications and significant at least 10%. We lose some of the significance in the MSA level regression due to the sample are cut in the middle of the housing booming (before 2003) while the magnitude remains (0.00431). The economic intuition of the result is that when house prices increase

Table 3: Summary Statistics III

This table reports the descriptive statistics for the sub-sample by low and high elasticity of states. The definition of variables are listed in Table A1. The sample is constructed from households in SIPP for the year 1996-2000, 2001-2004, 2004-2008 and 2008-2011. Low (High) elasticity is the bottom (top) tercile of the elasticity. A simple *t*-test using standard error clustered at state level is conducted to test the equality of the means between two groups. Standard errors clustered by state are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	Low Elasticiy	High Elasticiy	
	Mean	Mean	Difference
College Enrollment	0.6116	0.5749	0.0367*
			(0.0182)
Age	18.5410	18.5538	-0.0157
			(0.0204))
Female	0.5116	0.5169	-0.0053
			(0.0104)
Afrian American	0.1026	0.1589	-0.0564
			(0.0377)
Own	0.6765	0.7228	-0.0462
			(0.044)
Ln (1+Income)	8.2299	8.1304	0.0995
			(0.0626)
Ln (1+ Home equity)	7.0922	7.2741	-0.182
			(0.1988)
Ln (1 + Household non-housing wealth)	9.9654	9.9867	-0.0213
			(0.0551)

Table 4: College Enrollment by States

This table reports the summary statistic of college attendance ranked by elasticity. The definition of variables are listed in Table A1. The college attendance are the average between year 1996-2000 and 2001-2005, respectively.

			College Attendance					College Attendance	
		Elasticity	1996-2000	2001-2005			Elasticity	1996-2000	2001-2005
1	Utah	0.7500	0.5278	0.5051	20	DC	1.6100	0.5714	0.4167
2	New Hampshire	0.8600	0.7619	0.5682	21	Rhode Island	1.6100	0.5385	0.6364
3	Florida	0.8829	0.6116	0.5556	22	Virginia	1.7100	0.5698	0.6176
4	Wisconsin	1.0300	0.6575	0.5397	23	Tennessee	1.8067	0.5179	0.5241
5	Washington	1.0450	0.4762	0.5138	24	Michigan	1.9733	0.5593	0.6322
6	Oregon	1.0700	0.3333	0.4885	25	Alabama	2.0900	0.4894	0.5086
7	California	1.0727	0.6778	0.5914	26	New Mexico	2.1100	0.7500	0.5676
8	Maryland	1.2300	0.6563	0.6947	27	South Carolina	2.1833	0.5645	0.6525
9	Connecticut	1.2400	0.7353	0.7742	28	Kentucky	2.3400	0.6923	0.5287
10	Louisiana	1.2750	0.5455	0.6519	29	Ohio	2.4757	0.5904	0.4444
11	Nevada	1.3900	0.6667	0.4889	30	Georgia	2.5500	0.4384	0.5314
12	Minnesota	1.4500	0.6162	0.6980	31	Texas	2.7557	0.6514	0.5767
13	Arizona	1.5150	0.5965	0.4742	32	North Carolina	2.7667	0.4535	0.5510
14	Massachusetts	1.5200	0.7170	0.6784	33	Arkansas	2.7900	0.3810	0.5000
15	New Jersey	1.5600	0.7368	0.7455	34	Missouri	3.1900	0.5775	0.5615
16	Pennsylvania	1.5775	0.6000	0.5815	35	Oklahoma	3.3200	0.5345	0.6033
17	New York	1.5800	0.7080	0.6813	36	Nebraska	3.4700	0.6429	0.7429
18	Illinois	1.5850	0.6957	0.6285	37	Indiana	3.7000	0.6744	0.6372
19	Colorado	1.6000	0.6111	0.5732	38	Kansas	5.4500	0.5161	0.5726
						Mean	1.9509	0.5959	0.5852

Table 5: College Enrollment and Home Equity: Probit Regression

This table reports the baseline probit regression for the sample of college enrollment. The definition of variables are listed in Table A1. The sample is constructed from households in SIPP for the year 1996-2000, 2001-2004, 2004-2008 and 2008-2011. Column (3) reports the regression within sample that has MSA information before year 2003. Standard errors clustered by state or MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)
	Owners and Renters			Owners	
Ln (1+Home equity)	0.00435***	0.00441**	0.00431*	0.00521**	0.00702**
	(0.00162)	(0.00213)	(0.00246)	(0.00209)	(0.00289)
Ln (1+Income)	0.246***	0.247***	0.220***	0.217***	0.219***
	(0.0184)	(0.0189)	(0.020)	(0.0159)	(0.0167)
Female	0.247***	0.246***	0.252***	0.317***	0.310***
	(0.0213)	(0.0229)	(0.022)	(0.0265)	(0.0271)
African American	-0.117***	-0.109***	-0.150*	-0.113*	-0.0968
	(0.0392)	(0.0423)	(0.091)	(0.0588)	(0.0618)
Ln (1 + Household non-housing wealth)		-0.000545		-0.00555	
		(0.00628)		(0.00833)	
Marginal effect of increasing home equity	0.0017	0.0017	0.0012	0.0019	0.0025
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
State(MSA) Fixed Effects	Yes	Yes	Yes	Yes	Yes
Observations	15,199	14,311	7,148	10,758	10,148
Pseudo R-squared	0.0431	0.0432	0.0456	0.0410	0.0416

and raise the value of home properties, households can now extract more housing wealth from the house such that college-aged youths from the family are more likely to attend college. The effect is also economically significant. For instance, the specification with individual and financial controls and year and state fixed effects for home owners (column (5)). A one standard deviation increases of housing prices (namely a 44% increase from the average) implies that the average home property value ($131,614) increases by roughly $60,000 or 80% increase in home equity. The estimate of the marginal effect suggests that for an college-aged individual with the sample mean demographic and family income characteristics, and an 80% increase in home equity raises the probability that the individual attends college by $0.80 \times 0.0025 = 0.2$ percentage points. In terms of expected numbers, it is approximately 72,000 more students enrolled in college each year, which is not a negligible number.[1]

I also find that individual are more likely to attend college if they are female and if they are not black and if they come from family with higher income. The effect of non-housing wealth is however not significant, it is implying that other types of wealth do not seem to be very important in determining one individual's educational investment decision.

1.5.1.2 Robustness Checks

The baseline regression suggests that the appreciation of home equity would help individuals to overcome the credit constraint. However this can be attributed to omitted variables. In this section, I conduct a couple of robustness checks to further understand the underlying mechanism. First, I test whether the home equity borrowing channel of housing is stronger for individuals that are more credit constrained. The hypothesis

[1] According to the U.S. Census Bureau, there were 28.75 million young adults ages 18 to 24 in 2008.

Table 6: College Enrollment and Home Equity: Robustness Checks

This table reports the baseline probit regression for the sample of college enrollment. The definition of variables are listed in Table A1. The sample is constructed from households in SIPP for the year 1996-2000, 2001-2004, 2004-2008 and 2008-2011. Standard errors clustered by state are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level. Test "High = Low" is a *t*-test of equality of the *Ln(1+Home equity)* coefficients between high and low labor income.

	(1)	(2)	(3)	(4)
	Constrained	Unconstrained	2001-2005	Other years
Ln (1+Home equity)	0.00689**	0.00318	0.00825***	0.000110
	(0.00290)	(0.00412)	(0.00222)	(0.00221)
Ln (1+Income)	0.00110	0.686***	0.250***	0.242***
	(0.0336)	(0.117)	(0.0247)	(0.0218)
Female	0.237***	0.298***	0.233***	0.263***
	(0.0341)	(0.0432)	(0.0306)	(0.0298)
African American	-0.0702	-0.149**	-0.160**	-0.0770**
	(0.0629)	(0.0681)	(0.0643)	(0.0343)
Year Fixed Effects	Yes	Yes	Yes	Yes
State Fixed Effects	Yes	Yes	Yes	Yes
Test "High = Low"		1.07	3.675***	
Observations	5,077	5,058	8,114	7,080
Pseudo R-squared	0.0304	0.0416	0.0443	0.0492

is as follows: if the housing wealth has a positive effect for individual's college attendance decision, it must be more significant for individuals who are from families that are more financially constrained. I split the sample into labor income terciles using the level of monthly income at current year. The first tercile individuals have the lowest family income (constrained) while the third tercile individuals have the highest income (unconstrained). Column (1) and (2) in Table 6 present the results: the coefficient of home equity is positive and significant for individual from the family with low labor income (column(1)) while it is insignificant for those with high labor income (column (2)). This evidence suggests that housing wealth is more important in determining college attendance for credit constrained households.

Second, I continue to examine whether the effect of an increase in housing wealth varies with the business cycle. More specifically, the early-to-mid 2000s witnessed an exponential expansion in credit supply in general and in home equity borrowing in particular (Mian and Sufi (2009)), while home equity borrowing may not have been so prevalent in earlier decades. I repeat the baseline probit regression in the subsample between year 2001 and 2005, and also the subsample which contains the rest of years. The result shows a strong positive effect of housing wealth on college attendance during year 2001-2005 while no effect in the other periods. The difference between the two coefficients are statistically significant as well. It lends a strong support that during housing booming years, the housing collateral channel appears to have a much stronger effect on college attendance.

1.5.2 Instrumental Variables Results

The results of first-stage regression which are direct estimation of Equation (2) are presented in Table 7 column (1) and (2), I employ the measure of local housing elasticity from Saiz (2010). In those two columns,

I report the estimates from the first stage regression of the natural logarithm of home equity on the instrumental variable. The reported coefficients imply that the interaction of the change in national house prices between the year of purchase and the current year and housing supply elasticity, predict a large share of the variation in individual housing wealth. One can tell from the coefficient in interest Equation (2) is positive and significant at the 1 percent level. And also the value of the first-stage Wald Statistics (F-test), is strictly higher than the critical value for the IV regression to have no more than 10% of the bias of the probit estimate.

I next report the estimates from the second stage regression in Table 7. I use the the predicted home equity value from estimation of Equation (2) as an independent variable in Equation (1). Column (3) -(8) in Table 7 presents the instrument variable regression results when the home equity are instrumented by the housing supply elasticity and housing price. The coefficients estimated from this IV regression is larger than the results we obtain obtained from the baseline probit model. For example, the coefficients in Column (3) is positively significant at 1% confidential level. The estimates in Column (3) strongly suggests that the null hypothesis that current housing wealth has no effect on the college attendance is rejected. We can also find a similar pattern that the positive home equity effect is bigger for home owners alone. In Column (5) and (8) we use the sub-sample which contains the MSA-level geographic information to repeat the both baseline and IV regressions. And again, the results seem to support the findings using the state-level data. When it comes to economics significance, one can see that the coefficient in Column (3) is almost three times larger than the baseline probit regression. Again, marginal effect suggests that an college-aged individual with the sample mean demographic and family income characteristics, and an 80% increase in home equity raises the probability that the individual

Table 7: College Enrollment and Home Equity: IV Regression

This table reports the second stage of instrument variables regression. The definition of variables are listed in Table A1. In column (1) and (2), we report the first-stage OLS regression. In column (5) and (8), we use the sub-sample which contains the MSA-level geographic information to repeat the IV regressions (only the second stage is shown). In column (9), we report the reduced form regression which incorporates the instrument into the baseline regression. Standard errors clustered by state or MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	First stage (OLS)		Two-Stage Least Squares						Reduced form
			All (state)	All (state)	All (msa)	Owner (state)	Owner (state)	Owner (msa)	All (state)
House price growth times Housing supply elasticity	15.586***	13.688***							0.355
	(3.84)	(3.55)							(0.365)
Ln (1+Home equity)			0.012***	0.011***	0.015***	0.018***	0.017***	0.022***	0.013*
			(0.003)	(0.002)	(0.001)	(0.002)	(0.003)	(0.004)	(0.007)
Ln (1+Income)		0.245***	0.221***	0.223***	0.251***	0.245***	0.222***	0.221***	0.251***
		(0.018)	(0.017)	(0.017)	(0.025)	(0.023)	(0.015)	(0.013)	(0.016)
Female		0.243***	0.317***	0.298***	0.256***	0.248***	0.321***	0.311***	0.322***
		(0.025)	(0.03)	(0.026)	(0.023)	(0.025)	(0.027)	(0.028)	(0.04)
African American		-0.112***	-0.101*	-0.088	-0.118***	-0.109**	-0.115**	-0.111	-0.121**
		(0.042)	(0.052)	(0.062)	(0.039)	(0.043)	(0.055)	(0.064)	(0.052)
Ln (1 + Household non-housing wealth)		-0.001		-0.006		-0.002		-0.009	
		(0.008)		(0.006)		(0.007)		(0.009)	
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State (MSA) Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-statistics	121.5	56.8							
Marginal effect of increasing home equity			0.0126	0.0118	0.0135	0.0162	0.0172	0.0165	
Hausman test			2.58	3.12	2.01	3.12	2.88		
Observations	15,128	14,249	14,238	13,407	3,733	10,069	9,505	2,636	14,249
R-squared	0.0435	0.0424	0.0433	0.0423	0.0435	0.0447	0.0388	0.0411	0.0452

attends college by $0.80 \times 0.0126 = 1$ percentage points, which leads to approximately 289,800 more students enrolled in college each year.

Regarding the validity of the instrument, one can already observe in Column (1) and (2) that the instrument is strongly related to home equity. For the IV estimate to be consistent, it should be uncorrelated with the residual from the second stage specification. If the instrument is correlated with factors other than changes in individual housing wealth, the approach is invalid. In order to address this concern, we note that if the only impact of the instrument is through changes in housing wealth, then the instrument should be insignificant if included in Equation (1). Column (9) reports the estimate from this reduced form. One can tell from the insignificant coefficient that the test failed to reject the null hypothesis that the effect of changes in housing prices, times by the housing supply elasticity, is zero.

Table 7 also reports the result from a Hausman test which allows us to compare the OLS and IV regression results. As we can see in Tables 5 and 7, the 2SLS coefficient estimates are much larger than the corresponding OLS estimates. The test statistics from Hausman tests show that we can reject that the null hypothesis that IV and OLS models produce the same estimates.

In Table 8, we repeat the second-stage regressions by adding the state-specific trend. It is defined as the interaction of year and state. The results remain the same as in Table 7 that the coefficients for home equity are positively significant in all of the specifications.

1.6 Conclusion and Policy Implication

This paper shows that housing wealth alleviates credit constraints for potential college attendees by enabling home owners to extract equity from their property and invest it in the education. Using a large US individual-level survey dataset over the 1996-2011 period, I find that

one standard deviation increases of housing prices translate into approximately 72,000 more students enrolled in college each year. The effect is robust after controlling for a wide range of demographic and income characteristics and for location $\times$ year fixed effect. Given the large fluctuations in the housing market over the previous decade, including an unprecedented boom followed by a precipitous and sustained decline since 2006, identifying the effect of housing wealth on household decisions is of substantial policy interest.

This analysis contributes to the growing work indicating the importance of housing wealth for various types of household behaviors by separating the credit effect of housing wealth using a credible identification. My paper distinguishes from previous studies in the following ways: first, using interaction of land supply elasticity and interest rates as an instrument on housing prices captures the effect of credit in a better way. Second, the data used in my paper has a longer time span from 1996-2011 which covers both housing boom and bust. It provides a strong support on the effect of home equity on college attendance decision. My paper adds to the work of Corradin and Popov (2015) which shows the importance of credit constraints for new business creation. While the recent paper Sun and Yannelis (2015) also provides evidence on weather credit constraints exist and how they affect a household's college enrollment decision under the U.S. banking deregulation setup. My paper also complements to the growing literature on examining the collateral channel that helps relieving the credit constraints. It has been argued as being of first-order importance to firm investment (Chaney et al.(2011)), firm capital structure (Cvijanovic (2014)), entrepreneurship (Corradin and Popov (2015)) and innovation (Cao et al.(2014)).

The results of the paper have particular relevance to current policy as credit markets have tightened and housing prices have declined in many areas of the country. Considering the reduction in family resources caused by these problems in the housing market, it is likely that

many families will face increasing constraints in their ability to finance college in the near future. This work provides new evidence on that college attendance is sensitive to these fluctuations.

Table 8: College Enrollment and Home Equity: Robustness

This table reports the second stage of instrument variables regression including state-specific or MSA-specific trend. The trend is defined as the interaction of year and state or MSA. The definition of variables are listed in Table A1. In column (2) and (5), we report the regressions within the subsample that has the MSA level information before 2003. Standard errors clustered by state or MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)
	Probit	Probit	IV	IV
	Owner and Renter	Owner	Owner and Renter	Owner
Ln (1+Home equity)	0.005***	0.005**	0.015***	0.014**
	(0.002)	(0.002)	(0.002)	(0.002)
Ln (1+Income)	0.261***	0.235***	0.255***	0.249***
	(0.025)	(0.033)	(0.032)	(0.019)
Female	0.243***	0.312***	0.351***	0.317***
	(0.027)	(0.029)	(0.033)	(0.039)
African American	-0.142***	-0.108*	-0.113***	-0.121*
	(0.042)	(0.045)	(0.041)	(0.065)
Year Fixed Effects	Yes	Yes	Yes	Yes
State Fixed Effects	Yes	Yes	Yes	Yes
Trend	Yes	Yes	Yes	Yes
Observations	15,080	10,557	14,238	9,505
R-squared	0.0672	0.0711	0.0655	0.0691

Appendix I: Table A1

Variable Definitions

Variables	Definition
Key Variable	
College Attendance	A dummy variable which equals to one if the individual who graduated from high school and is currently enrolled in college
Demographics	
Age	The individual's age as of last birthday
Female	A dummy variable which equals to one if individual is female
African American	A dummy variable which equals to one if individual is African American
Housing and Wealth	
Own	A dummy variable equals one if the household owns the property they live
Home Equity	It denotes the difference between the value of the household's property and the value of the household's mortgage

Continued on next page

Table A1 - *Continued from previous page*

Variables	Definition
Income	It denotes the monthly income of the household
Household Non-housing Wealth	It denotes the total wealth of the household net of the amount of home equity
Home Property Value	It denotes the value of the household's property
Housing Prices	
Elasticity	Local Housing Supply Elasticity which comes from Saiz (2010)
State Residential Housing Prices	It denotes the State OFHEO real estate price index.
Mortgage Interest Rate	It denotes the "contract rate on 30-year, fixed rate conventional home mortgage commitments" from the Federal Reserve website

Table A2: College Enrollment by MSA

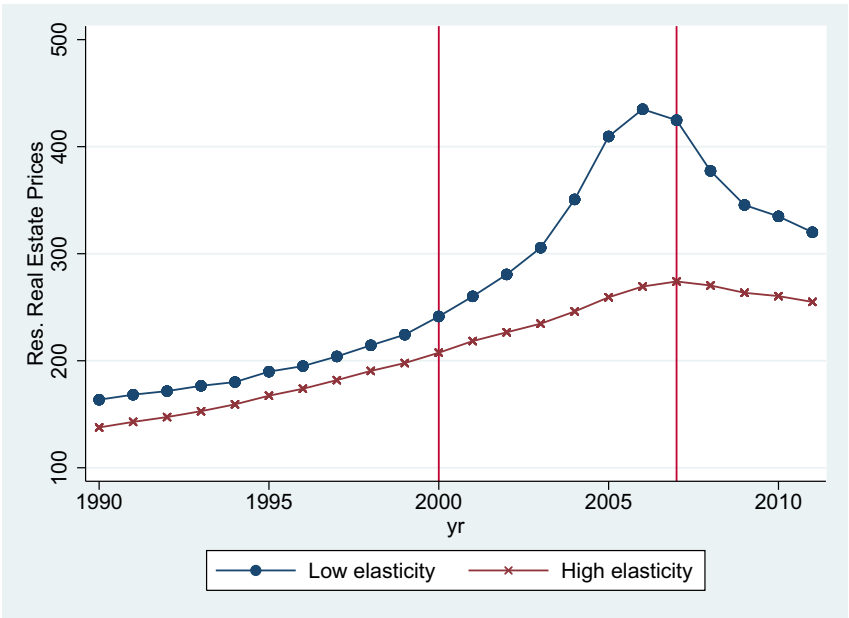
This table reports the summary statistic of top 10 and bottom 10 college attendance ranked by elasticity of MSA. The definition of variables are listed in Table A1. The college attendance are the average between year 1996-2011.

MSA	Elasticity	College Attendance	MSA	Elasticity	College Attendance
Miami-Fort Laudersale, FL	0.6250	0.6023	Dallas-Fort Worth, TX	2.4900	0.6698
San Diego, CA	0.6700	0.6818	Atlanta, GA	2.5500	0.5904
San Francisco-Oakland-San Jose, CA	0.7067	0.6667	Richmond-Petersburg, VA	2.6000	0.5769
Provo-Orem, UT	0.7500	0.4412	Columbus, OH	2.7100	0.3846
Salt Lake City-Ogden, UT	0.7500	0.5932	Greensville-Spartansburg-Anderson, SC	2.7100	0.4706
NY-NJ-LI, NJ	0.7600	0.7241	Killeen-Temple, TX	2.7557	0.5385
Los Angeles-Riverside-Orange County, CA	0.7850	0.6552	Corpus Christi, TX	2.7557	0.3636
New Orleans, LA	0.8100	0.6552	Beaumont-Port Arthur, TX	2.7557	0.6000
Norfolk-Virginia Beach-Newport News, VA	0.8200	0.6571	Fayetville, NC	2.7667	0.7778
West Palm Beach-Boca Raton, FL	0.8300	0.7647	San Antonio, TX	2.9800	0.5714
Boston-Worcester-Lawrence, MA-NH	0.8600	0.7130	Austin-San Marcos, TX	3.0000	0.5185
Melbourne-Titusville-Palm Bay, FL	0.8829	0.3333	Charlotte-Gastonia-Rock Hill, NC	3.0900	0.4706
Fort Pierce-Port St. Lucie, FL	0.8829	0.5000	Greensboro–Winston-Salem–High Point, NC	3.1000	0.4074
Pensacola, FL	0.8829	1.0000	Kansas City, KS	3.2900	0.8333
Fort Myers-Cape Coral, FL	0.8829	0.6667	Oklahoma City, OK	3.2900	0.6207
Lakeland-Winter Haven, FL	0.8829	0.3333	Tulsa, OK	3.3500	0.6296
Daytona Beach, FL	0.8829	1.0000	McAllen-Edinburg-Mission, TX	3.6800	0.6667
Sarasota-Bradenton, FL	0.9200	0.5000	Dayton-Springfield, OH	3.7100	0.7742
Tampa-St. Petersburg-Clearwater, FL	1.0000	0.5926	Indianapolis, IN	4.0000	0.6829
Madison, WI	1.0300	0.4545	Fort Wayne, IN	5.3600	0.5000

Figure 1: Evolution of State Residential Prices (High versus Low Elasticity, 1900-2011)

This figure shows the average state residential real estate prices (normalized to 100 in 1980) for States in the bottom tercile of land supply elasticity (“Low Elasticity") in a dot line and States in the top tercile of land supply elasticity (“High Elasticity") in a X line.

Figure 1.1



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2. Income Inequality and Trade

Does trade with unskilled labor-abundant countries reduce the relative wages of U.S. unskilled labor and consequently cause increased income inequality across industries and regions? Empirical studies in the 1990s found only a modest effect. In this paper, I re-consider the question by using the income inequality measures constructed from Current Population Survey (CPS) data and analyzing the effect of rising Chinese import competition between 1993 and 2007 on US local labor markets. I find that areas which are more exposed to China imports competition have larger changes in income inequality. In my main specification, a $1,000 exogenous decadal rise in a MSA's import exposure per worker leads to a 1.5% increase in the logistic Gini. This redistributive effect is more profound among non-college educated workers in manufacturing sectors.

2.1 Introduction

It has been a major topic of research in labour economics over the last two decades to understand changes in the wage distribution. Beginning in the 1990s, the literature developed rapidly as economists sought to understand the determinants behind the rising wage inequality, particularly in developed countries like the US. One of the measures of income inequality, the relative wages of college graduate workers to high school graduates, has shown a tendency to increase over multiple decades, despite the large secular increase in the relative supply of college students. According to Van Reenen (2011), although US wage inequality rose monotonically in the 1980s, since the early 1990s a better description would be “polarization".

The importance of understanding the causality of income inequality has been emphasized both in academics and in practice. Inequality is among one of the central of political debates nowadays: there is a sense that President Obama put it mildly when he declared inequality “the defining challenge of our generation." Joseph Stiglitz has written that, “with inequality at its highest level since before the Depression, a robust recovery will be difficult in the short term, and the American dream – a good life in exchange for hard work – is slowly dying." Paul Krugman calls inequality “toxic."

One of the biggest concerns for rising income inequality is that it might be eroding the equal opportunity and mobility of the entire social system. As the rungs of the economic ladder move further and further apart, conventional wisdom says that it will become much more difficult to climb them. Opportunities for upward mobility will disappear as the deck becomes stacked against the middle class and the poor. Moreover, countries with lots of inequality tend to develop unhealthy political systems, and eventually to fail. As Acemoglu and Robinson write in their book “*Why Nations Fail*" that “...But the real reason to worry about [economic inequality] is not the unfairness of it all... The problem is that

economic inequality often comes bundled with inequality of opportunity and political inequality..." They point out that economic inequality will lead to greater political inequality, and those who are further empowered politically will use this to gain a greater economic advantage by stacking the cards in their favor and increasing economic inequality yet further - “a quintessential vicious circle."

It is however, still an open question regarding the causes of income inequality in the U.S. While standard economic analysis predicts that increased US trade with unskilled labor-abundant countries should reduce the relative wages of U.S. unskilled labor, trade in the form of foreign outsourcing was associated with modest increases in the wage premium for skilled manufacturing labor (Feenstra and Hanson, 1999). The current empirical evidence seem to suggest that other shocks, including skill biased technical change, played a more important role in the evolution of the U.S. wage structure in the last couple of decades (Van Reenen (2011)). The economists were reaching the contemporaneous consensus that trade had not in recent decades been a major contributor to declining manufacturing employment or rising wage inequality in developed countries.

Just as the consensus on the consequences of trade for wages and employment has been reached among economists, an epochal shift in patterns of world trade was gaining momentum. China, which has been lagged behind for centuries, was eventually emerging as a great power, and toppling established patterns of trade accordingly. In 1991, low-income countries accounted for just 9 percent of US manufacturing imports (Autor, Dorn and Hanson (2013)). In 2000, the low income-country share of US imports reached 15 percent and climbed to 28 percent by 2007, with China accounting for 89 percent of this growth. The share of total US spending on Chinese goods rose from 0.6 percent in 1991 to 4.6 percent in 2007, with an inflection point in 2001 when China joined the World Trade Organization (WTO). The consensus that trade

could be strongly re-distributive in theory but was relatively benign in practice has not stood up well to these new developments.

In this paper, I re-examine the relationship between changes in income inequality across US local labor market and change in exposure to Chinese import competition between 1993 and 2007. The time span covered in this study enables me to take the surge of trade between US and developing countries like China into account. Since the trade shocks work through in general equilibrium, one needs empirically to map many industry-specific shocks into a small number of aggregate outcomes. Focusing only on national labor market at annual frequencies leaves one with few observations and many confounding effects. To overcome the degree-of-freedom problems in estimating the effects of trade shocks, following prior studies, I treat the industry or occupation as the unit of analysis. This approach to analyzing the impact of trade exposure follows important recent work by Liu and Trefler (2008), who estimate that over 1996-2006, US outsourcing of services to China and India had minimal effects on changes in occupation, employment, or earnings for US workers. Ebenstein et al. (2013), who use data from the CPS, find larger effects of trade on wages, with wages growing more slowly in occupations more exposed to import penetration and to US multinationals moving production offshore.

An alternative solution to the degree-of-freedom problem to take regional economics as the unit of analysis. I relate changes in exposure to low-income-country imports to changes in MSA-level measures of income inequality, where I allow impacts to vary by age, gender, and education. The local market approach is aliened with recent work by Chiquiar (2008), Topalova (2005, 2010), and Kovak (2013), who study the effects of trade liberalizations on wage, poverty, and migration in local and regional labor markets in developing countries in Mexico, India, and Brazil, respectively.

As mentioned in Krugman (2008), the lacking of up-to-date the data

are always being rarely mentioned despite of the reassuring consensus that trade has only modest effects on income distribution. Although there have been a lot of later developments, it is still an open question. Autor, Dorn and Hanson (2013) find that local labor markets that are exposed to rising low-income-country imports due to China's rising competitiveness experience increased unemployment, decreased labor-force participation. While Liu and Trefler (2008) find negligible effects on changes of labor market outcomes at industrial and occupational level. And Ebenstein et al. (2013) find larger wage effects in occupational exposure to trade penetration from low-income-country than industry exposure. In this paper, I provide some evidence on the effect of trade on direct measures of income distribution. Focusing on the direct measures of income distribution such as college/high-school wage premium, Gini coefficient, Theil index instead of wage level, employment participation, occupation and industry switching distinguish my paper from prior research.

China's rise provides a natural opportunity to study impact of a large trade shock on labor markets in developed countries. My paper belongs to a growing literature that offers rich evidence and surprises that should catalyze research for many years to come. By revisiting the relationship between change in income inequality and trade, the contribution of this paper is at least two folds. Firstly, I believe that this evidence calls into question the consensus of the early 2000s, and reinforce that, after the early Bretton Woods era aberration, the distributional consequences of trade are effective. More importantly, it shows that the free trade among countries could have both benefits and costs. And the costs, especially trade adjustments costs, have only been recognized in theory and was not empirically evident. Thus a better understanding when and where trade is costly, and how and why it may be beneficial, are key items on the research agenda for trade and labor economists. Secondly but more importantly, it is crucial to develop effective tools for managing and mit-

igating costs of trade adjustment.

Most of the empirical studies, however, generally focused on the absolute change of wage or employment, namely the change of level of number of workers or change of average wage. In this paper, I rely on the identification restriction suggested by Autor, Dorn and Hanson (2013) to measure China's shock; namely, I use the predicted changes in U.S. imports from China using as instrument the change in imports from China by other high-income countries for the period 1993 to 2007. My paper complement the prior studies in the following ways: first and most significantly, I look at the re-distributive effect of international trade among the entire income distribution rather than solely focusing on the absolute level. Secondly and more importantly, I measure the trade exposure at MSA level rather than commuting zone (CZ) level in Autor, Dorn and Hanson (2013). Commuting zones (CZs), which encompass all metropolitan and non-metropolitan areas in the United States, are logical geographic units for defining local labor markets. Being unable to differentiate between metropolitan and non-metropolitan areas potentially affects the estimates of wage effects in Autor, Dorn and Hanson (2013). They find a positive, but insignificant impact of import competition on local wages, leading them to conclude that "manufacturing plants react to import competition by accelerating technological and organizational innovations that increase productivity and may raise wages". My research suggests that the re-distributive effect of trade on local wages is at least significant at MSA level because the movement of workers out of manufacturing and into lower wage services happens more likely within MSA (Ebenstein et al. (2013)), which puts significant downward pressure on wages and thus widening the income distribution. Also closely related to me is Campbell and Lusher (2016). They find that workers in industries with greater exposure to international trade were more likely to be unemployed a year later and also experiencing a wage cut. Meanwhile, their findings do not suggest that

import exposure leads to increase of income inequality. There are two factor could possibly lead to the different outcomes between their paper and mine: first, the datasets are structurally different. Their analysis that following Ebenstein et al. (2012) is solely based on sectors while my paper has both regional and sector layers. Second, they use variation in exchange rate as instrument for imports at sectoral level which is different from what I use in this paper.

My paper complements the strand of literature that estimates the effects of trade shocks by treating the industry or occupation as the unit of analysis. This approach to analyzing the impact of trade exposure follows important recent work by Liu and Trefler (2008), who estimate that over 1996-2006, US outsourcing of services to China and India had minimal effects on changes in occupation, employment, or earnings for US workers. Ebenstein et al. (2013), who use data from the CPS, find larger effects of trade on wages, with wages growing more slowly in occupations more exposed to import penetration and to US multinationals moving production offshore. In examining economic outcomes at the level of local labor markets, I am able to capture both the direct effect of trade shocks on income distribution at import-competing employers as well as the net effects.

2.2 Literature Review

Stemmed from the standard textbook analysis which says that the trade is driven by international differences in factor abundance, the classic Stolper and Samuelson (1941) theorem implies trade in theory can have very strong effects on income distribution. In particular, if trade with unskilled labor abundance countries leads to a reduction in the relative price of unskilled labor-intensive good should reduce the real wages of less educated workers, both relative to other workers and in absolute terms. The more recent representative studies focusing on general

equilibrium effects of trade on wages include Baldwin and Cain (2000), Baldwin and Hilton (1984), Krugman (2000), and Leamer (1994, 1996, 2000). Many of these studies use a specific formulation of the Stolper-Samuelson theorem which usually contain a two-good, two-factor, and two-country framework to illustrate the relationship between trade and factor prices.

But is the effect of trade on wages quantitatively important? There is considerable disagreement among economist concerning the role of international trade in accounting for the income distribution. Deardorff and Hakura (1994) review seven empirical studies and report that findings range from trade having a major impact on the distribution of wages to having no impact. There are five additional studies not included in this review are: Berman, Bound and Griliches (1994), who conclude that shifts in trading patterns contributed to widening the observed inequality in relative wages but that other forces, particular biased technological process in favor of more skilled workers, had a larger impact; Borjas and Ramey (1994), who find that increased international competition explains “much of the increase in wage inequality,"; Krugman and Lawrence (1993), who argue that the effect of increased international competition on the widening of U.S. income inequality has been “quantitatively minor"; and Leamer (1996), who concludes that “a very large percentage of the increased income inequality among workers can be attributed to globalization, but most of this effect came during the 1970s"; and Sachs and Shatz (1994, p.57) who conclude that ”increased trade has contributed...to the growing inequality of earnings between low-skilled and high-skilled workers, although the weight of the trade effect is uncertain."

Because of the decrease in concerns about wages during the economic upturns in the later 1990, the discussion in international economics then slided away from the distributional effects of trade exposure with developing countries. Nevertheless, employing the most up-to-date

data in the 21 century, recent studies still find the relationship between trade and income inequality ambiguous. Edwards and Lawrence (2010) or Ebenstein et al. (2013), which observe no negative wage effects of imports on US workers in import-competing manufacturing industries. Bloom, Draca, and Van Reenen (2011) argue that manufacturing plants react to import competition by accelerating technological and organizational innovations, in such way, one would not be able to observe substantial wage effects of imports on worker's wage. And lastly, Autor, Dorn and Hanson (2013) and Acemoglu et al. (2014), both find that increase in U.S. imports from China, which accelerated after 2000, was a major force behind recent reductions in U.S. manufacturing employment at both industry-level and local labor market-level. In most of these empirical studies, the main reason U.S. labor markets are differentially exposed to Chinese goods is their different degree of import competition.

Representative studies focusing on general equilibrium effects of trade on wages include Baldwin and Cain (1997), Baldwin and Hilton (1984), Krugman (2000), and Leamer (1994, 1998, 2000). Most of these studies begin with a simple illustration of the relationship between trade and factor prices using specific formulation of the Stolper-Samuelson theorem with a two-good, two-factor, and two-country framework. For example, Leamer (1994) begins with a typical zero profit condition that can be written in vector form as $Aw = p$, where p is a vector of product prices, w is the vector of factor costs and A is the vector of input intensities. If skilled and unskilled labor are the only two factors and texiles (T) and machinery (M) are the only two goods, then Leamer reproduces the standard general equilibrium result using the Heckscher-Ohlin framework that wage and income inequality will depend on both technology (as represented by the A matrix) and product prices.

More specifically, if we assume that textiles are unskilled labor intensive and machines are skilled labor intensive, then the wage of un-

skilled workers will vary positively with p_T and negatively with p_M; this results is reversed for skilled workers, i.e. if the price of skill labor-intensive product p_M rises, it must lead to a rise in the relative wages of skilled workers. This implies a fall in the ratio of skilled to unskilled workers in both industries - and hence a fall in the marginal productivity of unskilled workers in terms of both goods. And that, in turn, means that the real wage of unskilled workers unambiguously falls. This is the theoretic base for re-distributive effects of international trade.

What does it look like when one applies the Stolper-Samuelson theorem on the current development of trade between the U.S. and China? Consider first the change of factor prices. The BLS publishes price indices of manufactured imports from developing and developed countries. If we assume that developing countries mainly export unskilled labor-intensive goods to the United States whereas developed countries export skilled labor-intensive goods, the ratio of these prices should offer a measure of the relative price of unskilled labor-intensive goods. Figure 2 shows the logarithm of this ratio, normalized so that 1995 =0. It indeed seems to have been a substantial decline in the relative price of unskilled labor-intensive goods since the mid-1990s.

Following Krugman (2008), I move on to analyze changes in income distributions. Figure 3 shows two indicators of wage differentials commonly used in practice: the 90-50 ratio of hourly wages and the college-noncollege ratio. Both are shown for men only, to abstract from changes in sex differentials; both are also expressed in logarithms, normalized so that 1995=0. One can see from Figure 3 that both measures have risen dramatically since 1995. Thus the combination of the price changes shows in Figure 2 and the income distribution changes shown in Figure 3 does look reasonably supportive of proposition that rapid growth in the exposure of import competition to developing countries has made the effects on inequality substantially larger. Figure 2 and 3 demonstrate the potential association between the rising of import from

developing countries and the increasing of income inequality. In order to establish causality between the two, we need more rigorous econometric setup and better source data.

2.3 Data source and Sample

Information on the distribution of income is from the March Supplement of the Current Population Survey (CPS), which is an annual survey of over 60,000 households across US. The CPS is a repeated, representative sampling of the population, but it does not trade individuals through time. For industry trade exposure data, I employ data from the UN Comrade Database on US imports at the six-digit Harmonized System (HS) product level. Following international trade literature, e.g. Bardhan and Kroll (2003), I link service-trade data to industries and occupations using standard concordances.

In Figure 1, I show the value of annual US imports from China from the years 1991-2007 (with all values in 2007 US$). During the 16 year period from 1991 to 2007, this import value increased by a factor of 11.5, from $26 billion to $330 billion. Based on the UN Comrade Trade data, the volume of U.S. exports was significantly smaller than the volume of imports throughout these years, and the growth of imports outpaced the growth of exports. The primary change in US-China trade during our sample period is thus the dramatic increase of US imports. Figure 1 also presents the value of U.S. imports from Canada, Mexico and Central America[1] from year 1991 to 2007. One can observe that although imports from these countries also grew considerably overtime, it is however much less dramatic than in the case of Chinese imports. For example, imports from Canada started from a much higher level than Chinese imports. It has been surpassed in 2004. One thing also worth

[1]Mexico/CAFTA includes Mexico, the Dominican Republic, and all Central American countries except Belize and Panama.

noticing is that there is an inflection point in 2001 when China became a member of World Trade Organization (WTO).

To access the effect of imports of Chinese goods on local labor markets, I need to define regional economics in the US. I hereby employ the Metropolitan Statistical Area (MSA) as the concept for local labor markets. A recent analysis based on MSA level is Aghion et al. (2015) which analyze the relationship between turnover-driven growth and subjective well-being, using cross-sectional MSA level US data. The effects of Chinese imports will vary across local labor markets in the US because there is substantial geographic variation in industry specialization. Regions would respond more strongly where their local economies are specialized in industries which outputs compete with Chinese imports. The measure for the exposure of local labor markets to Chinese imports combines trade data with data on local industry employment. I explain how the data is constructed in Data Appendix.

Following Beck, Levine and Levkov (2010), I measure the distribution of income for each industry-year to begin with. First, the Gini coefficient of income distribution is derived from the Lorenz curve. The Gini coefficient equals zero if everyone receives the same income, and equals one if a single individual receives all the economy's income. I present results with both the natural logarithm of the Gini coefficient as well as the logistic transformation of the Gini coefficient in the regression analyses. The logarithmic transformation of the Gini coefficient removes the floor and makes the measure upper-bounded at zero. The advantage of using the log of Gini is that it allows us to interpret coefficients relative to this variable as percentage changes. My second measure of income distribution is the Theil index, which is also increasing in the degree of income inequality. Third, I examine the difference between the natural logarithm of incomes of those at the 90th percentile and those at the 10th percentile (Log(90/10)). Finally, I would use the difference between the natural logarithm of incomes at the 75th percentile and that at the 25th

Table 1: Summary Statistics

This table reports the descriptive statistics for variables used in the paper. I used five measures of income inequality: a logistic transformation of the Gini coefficient, the log of the Gini coefficient, the log of the Theil index, the log ratio of the 90th and 10th percentiles of the income distribution, and the log ratio of the 75th and 25th percentiles of the income distribution. We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. Data on real per capital GDP are obtained from the Bureau of Economic Analysis. Data on unemployment rate are from Bureau of Labor Statistics. The proportions of different employment types are calculated from CPS data.

	Mean	Median	SD	p25	p75
Income Inequality					
Logistic Gini coefficient	-0.28	-0.29	0.07	-0.35	-0.25
Log Gini coefficient	-0.85	-0.85	0.03	-0.87	-0.83
Log Theil index	-1.15	-1.17	0.08	-1.23	-1.11
Log 90/10 ratio	2.40	2.46	0.21	2.33	2.59
Log 75/25 ratio	1.15	1.14	0.11	1.09	1.25
Log 90/50 ratio	0.95	0.85	0.08	0.85	1.03
Demographics					
Growth rate of per capita GDP (2000 dollars)	0.02	0.02	0.03	0	0.025
Percentage of blacks	0.10	0.06	0.11	0.02	0.15
Percentage of college-educated	0.75	0.74	0.04	0.71	0.82
Percentage of employment	0.96	0.96	0.01	0.95	0.97
Percentage of self-employed	0.11	0.11	0.03	0.09	0.13
Percentage of skilled self-employed	0.07	0.07	0.02	0.03	0.05
Percentage of salaries workers	0.88	0.88	0.03	0.86	0.90
Unemployment rate	4.83	4.80	1.2	4.0	5.6

percentile ((Log(75/25))). Following the same procedure, I also construct the MSA-year income inequality measures. In Appendix table, I describe each measure as well as their advantages and disadvantages in details.

In total, there are 4650 observations (310 MSAs across 15 years). Table 1 presents the summary statistics for variables used in the paper. I use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. Data on real per capital GDP are obtained from the Bureau of Economic Analysis. Data on unemployment rate are from Bureau of Labor Statistics. The proportions of different employment types are calculated from CPS data.

2.4 Empirical Modeling

2.4.1 Instrument Variable Model

At region level regression-based analysis, I use the following econometric model to examine the relation between income inequality and trade exposure to China import competition:

$$\Delta Y_{it} = \beta_0 + \beta_1 \Delta IPW_{uit} + \beta_2 X_{it} + \nu_{it} \tag{2.1}$$

In equation (1), ΔY_{it} is a change of a measure of income distribution in region i in year t, X_{it} is a set of time-varying, regional level variables and ν_{it} is the error term. The variable of interest, ΔIPW_{uit} is the change in import exposure. Autor, Dorn and Hanson (2013) develops a general equilibrium model that provides a theoretic argument on how this reduced form can be employed to analyze the China imports effects on the wage and labor market of US region i. In short, the key assumption underlying the model is to assume trade has a “gravity" structure,

in which case one can map changes in trade quantities into labor-market outcomes.

I sketch the model as follows: they begin by describing how changes in China's export supply affect U.S. industry product demand, and subsequently transmitted to changes in traded sector output at the U.S. regional level. They assume the total demand by the U.S. aggregate economy for traded output produced in U.S. region i as,

$$M_i = \sum_j \frac{A_{ij}\tau_{ij}^{-\theta}}{\Phi_j} E_j \tag{2.2}$$

whereM_i is total sales by region i in the aggregate U.S. market, A_{ij} is the production capability of region i in industry j, τ_{ij} is the average transport cost for region i in shipping goods in industry j to the aggregate U.S. market, θ is the trade cost elasticity, E_j is the U.S. aggregate expenditure on goods in industry j, and Φ_j is the competitiveness of the aggregate U.S. market in industry j, defined as $\Phi_j \equiv \sum_{i'} A_{i'j}\tau_{i'j}^{-\theta}$

They move on to analyze the impact on traded output in region i of changes in production capabilities in the regions that supply the U.S. by totally differentiating (2), and using $\hat{m} \equiv dm/m$, they obtain the following equilibrium condition for the log change in region i's output:

$$\hat{M}_i = \sum_j \phi_{ij}\hat{E}_j - \theta\hat{\omega}_i + \sum_j \phi_{ij}\hat{A}_j - \sum_j \phi_{ij} \sum_{i' \neq c} \rho_{i'j}\hat{A}_{i'j} - \sum_j \phi_{ij}\rho_{cj}\hat{A}_{cj} \tag{2.3}$$

where $\phi_{ij} \equiv M_{ij}/M_i$ is the share of industry j in region i's total sales on the U.S. market, $\rho_{ij} \equiv M_{ij}/E_j$ is the share of region i in total U.S. purchase in industry j, and subscript c indexes China. And they assume that $\hat{A}_{ij} = \hat{A}_j - \theta\hat{\omega}_i$, and that trade costs remained unchanged.

Equation (3) provides a reduced-form specification for estimating the impact of trade shocks emanating from China on regional economic activity in the U.S. or other countries. Most empirical analyzes of the China trade shock base estimation on a specification like (3). The pri-

mary interest is the final terms on the right of (3), which presents the impact of growth in China's productive capacity on traded output by U.S. region i. If one plug in ρ_{ij}, it can be rewritten as,

$$\sum_j \phi_{ij}\rho_{cj}\hat{A}_{cj} = \sum_j \phi_{ij}[\frac{M_{cj}\hat{A}_{cj}}{E_j}] \tag{2.4}$$

which is the weighted average exposure of region i to changes in U.S. industry import penetration that is mandated by changes in China's production capacities. Equation (4) can be further simplified and rewritten as follows.

Following Autor, Dorn and Hanson (2013), the main measure of local labor market exposure to import competition is the change in Chinese import exposure per worker in a region, where imports are apportioned to the region (L_{ijt}) according to its share of national industry employment(L_{ujt}):

$$\Delta IPW_{uit} = \sum_j \frac{L_{ijt}}{L_{ujt}} \frac{\Delta C_{ucjt}}{L_{it}} \tag{2.5}$$

In equation (5), L_{it} is the start of period employment (year t) in MSA i and ΔC_{ucjt} is the change in US ($\boldsymbol{U}$) imports from China ($\boldsymbol{C}$) in industry j between start and end of the period.

As mentioned above, in order to identify the causal effect of rising Chinese import exposure on income distribution, they employ an IV strategy that accounts for the potential endogeneity of US trade exposure. The fact that rising competitiveness of Chinese manufacturers and China's lowering of trade barrier, dismantling of central planning, and accession to the WTO enable us to confidently implement the following instrument to identify the supply-driven component of Chinese imports. I instrument for growth in Chinese imports to the United States using the contemporaneous composition and growth of Chinese imports

in eight other developed countries.[1] Specifically, I instrument the measure import exposure variable ΔIPW_{uit} with a *non*-US exposure variable ΔIPW_{oit} which using data on contemporaneous industry-level growth of Chinese exports to other high-income markets:

$$\Delta IPW_{oit} = \sum_j \frac{L_{ijt-1}}{L_{ujt-1}} \frac{\Delta C_{ocjt}}{L_{it-1}} \tag{2.6}$$

This expression for *non*-US exposures to Chinese imports distinguish the expression in equation (5) in two respects. First, I replace the realized US imports by industry ΔC_{ucjt} with realized imports from China to other high-income markets (ΔC_{ocjt}). Second, according to Autor, Dorn and Hanson (2013), they replace the start-of-period employment levels by industry and region in equation (5) with employment levels from the prior decade and they rationale as follows: to the degree that contemporaneous employment by region is affected by anticipated China trade, the use of lagged employment to apportion predicted Chinese imports to regions will mitigate this simultaneity bias.

It is the same instrument variable approach suggested by Autor, Dorn and Hanson (2013), they employ an instrumental variable that accounts for the potential endogeneity of US trade exposure. The reason for the potential endogeneity problem is that Chinese import exposure is that some of the import is intermediate goods in the U.S. manufacturing sector. A general increased demand may therefore lead to an increase in import from Chinese but not necessarily increased competition from Chinese products. This instrument variable fulfills the two important properties: correlated with the endogenous independent variable and uncorrelated with the error term. First of all, it is obviously that Chinese imports in eight other developed countries are related to

[1]The eight other high-income countries are those that have comparable trade data covering the full sample period: Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland.

the imports to the U.S., I provide more evidence on Appendix Table 2. On the other hand, the rise of imports from China originates from the rising competitiveness of Chinese manufacturers themselves; in other words, the rise is a supply shock from the perspective of United States producers alongside with China's internal reforms, for instance, lowering of trade barriers, dismantling of central planning, and accession to the WTO. The momentum of the transition, which has propelled China's trade, is mostly due to dismal conditions in China at the time of Mao's death than to China's subsequent responses to contemporaneous shocks in high-income economics. Therefore, the second property of the instrument variable as of being 'valid', should hold in theory.

I use 2SLS, as I argued, the instrument variables should be in theory un-correlated with any unobserved residuals in the second stage regression, thus applying the predicted Chinese imports from the first stage regression to the second stage would relieve the following concern to some extent: the realized US imports from China and income distribution might both be positively related to the unobserved shocks to US product demand, in which case the OLS estimate of how increased imports from China may understate the true impact to the income inequality. Then it is possible to obtain a consistent estimation for the impact.

There are a couple of key features worth noticing about making China's rise as trade shocks. The first is the unexpected nature of China's export growth, which caught many observers by surprise. Not until recently two decades that few started foreseeing the importance of China for the world economy. Between 1984 and 1990, China's share of world manufacturing export increased modestly, from 1.2% to 1.9%. Its expansion did not begin until the 1990s. Secondly, China's deep isolation under Mao era created huge space for its later catch up. Between 1952 and 1978, China's GDP per capital increased from 59th in the world to 134th out of 167 Penn World Trade Table countries. By 1991, China's income ranking had raised up from 134th to 126th in the world. By

2001, China's income ranking had risen to 101st and by 2011 it had reached 77th.

Appendix Table 2 presents the operation of the instrumental variable strategy. The first column shows the annual imports from China increased by 304 billion dollars between 1991 and 2007, while imports from china grew by 235 billion dollars across the eight other high-income countries offering comparable trade data for the full sample period. Both the U.S. and the other high income countries experienced rising imports in almost all of hte 397 harmonized four-digit manufacturing industries, and the pattern of import growth across industries is highly correlated between the U.S. and the other countries (correlation coefficient of 0.92). The remaining columns of the table show the same information separately for each of the eight other high income countries. Each of the comparison countries witnessed import growth in at least 343 of the 397 manufacturing industries, and industry patterns of imports are strongly correlated with the U.S., with correlation coefficients ranging from 0.55 (Switzerland) to 0.96 (Australia). The fact that China made comparable gains in penetration by detailed sector across various countries in the same time interval suggests that China's falling prices, rising quality, and diminishing trade and tariff costs in these surging sectors are a root cause.

2.5 Empirical Results

I present the baseline results (Equation (1)) in Table 2. All regressions include geographic dummies for MSAs that absorb region-specific trends in the income inequality. A vector of MSA-level control variables are also included in the regression, for example, the percentage of blacks, the real growth rate of per capita GDP, the unemployment rate, the percentage of college-educated, and the percent of female-headed

households. I cluster standard errors at the MSA level.

I find modest evidence that exposure to Chinese imports competition is positively associated with income inequality during the sample period from 1993 to 2007. The estimated coefficients are statistically significant in two of five income inequality measures. Meanwhile, I also find that several of the economic and demographic MSA controls are significant predictors of income inequality. For instance, a higher unemployment rate, a higher proportion of blacks, and a lower proportion of college-educated are all positively associated with income inequality. In turn, a higher per capita GDP growth reduces inequality, but the effect is only significant in one specification.

Table 3 presents the results of instrumental variable regressions. The second stage lends strong support to that import exposure per worker is positively associated with income inequality and estimated coefficients are statistically significantly across all specifications. The coefficient of 0.015 in column 1 indicates that a $1,000 exogenous decadal rise in a MSA's import exposure per worker leads to a 1.5% increase in the logistic Gini. In other words, according to the 2013 world bank income inequality data, 1.5 % more in the logistic Gini represents the difference between Sweden and Afghanistan. To assess the economic relevance of this result, I compare the coefficient estimate to the demeaned standard deviation of the logistic Gini that I obtain after taking care for MSA and year effects. Since the standard deviation is 7% (see Table 1), the import exposures explain 21.4% of the variation in income inequality relative to region and year averages. First stage F-statistics indicates the instrument variable, the Chinese import exposure for other countries, are not a weak instrument. The t-statistics for Hausman tests indicate that the null hypothesis of IV and OLS models generate the same estimates can be rejected.

I repeat the same analysis with different age groups. Specifically, I use three additional age groups: 16-34, 35-49, and 50-64. For each case,

I compute the measures of inequality and then run similar regressions as in Table 3. Column (1) to (5) provide 2SLS second stage estimates while column (6) to (10) present OLS regression results. The results are presented in Table 4, one can see from column (1) to (5), the 2SLS second stage estimates stays statistically significant and the magnitudes remain similar as before. Meanwhile, the OLS results are insignificant across all sub-groups. In table 5, I further to analyze the re-distributive effect of China trade on different sectors and sex. In Panel II of Table 5, it shows that the effect of trade exposure on income inequality are similar for males and females. In Panel I, the effect is much more significant in manufacturing sectors. The effect is almost non-existed in non-manufacturing sectors like financial service, real estate or military.

Table 2: The Impacts of China Imports on Income Inequality: OLS

The table shows the impacts of imports from China on income inequality under standard OLS regression. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Gini coefficient, (3) the natural logarithm of the Theil index, (4) the natural logarithm of the ratio of the 90th and 10th percentiles, (5) the natural logarithm of the ratio of the 75th and 25th percentiles, and (6) the natural logarithm of the ratio of the 90th and 50th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25	Log 90/50
(Imports from China to US)/worker	0.009	0.005	0.01*	0.025*	0.008	0.002
	(0.006)	(0.003)	(0.006)	(0.015)	(0.005)	(0.007)
Percentage of blacks	0.545***	0.321***	0.665***	1.067**	0.614**	0.522**
	(0.194)	(0.112)	(0.224)	(0.485)	(0.285)	(0.251)
Growth rate of per capita GDP	-0.103	-0.059	-0.095	-0.178	-0.204*	-0.211
	(0.086)	(0.049)	(0.104)	(0.281)	(0.182)	(0.121)
Unemployment rate	0.011***	0.006***	0.013***	0.041***	0.014***	0.012*
	(0.003)	(0.002)	(0.004)	(0.011)	(0.005)	(0.065)
Percentage of college-educated	-0.630***	-0.395***	-0.760***	-0.843*	-0.501**	-0.425**
	(0.153)	(0.088)	(0.191)	(0.457)	(0.198)	(0.205)
MSA dummies	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.133	0.132	0.170	0.354	0.285	0.275

Table 3: The Impacts of China Imports on Income Inequality: 2SLS Estimates

The table shows the impacts of imports from China on income inequality. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Gini coefficient, (3) the natural logarithm of the Theil index, (4) the natural logarithm of the ratio of the 90th and 10th percentiles, (5) the natural logarithm of the ratio of the 75th and 25th percentiles ,and (6) the natural logarithm of the ratio of the 90th and 50th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. First stage estimates in panel II also include the control variables that are indicated in the corresponding columns of panel I. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25	Log 90/50
	2SLS second stage estimates					
(Imports from China to US)/worker	0.015**	0.008***	0.025***	0.038**	0.01**	0.015*
	(0.006)	(0.003)	(0.006)	(0.015)	(0.005)	(0.008)
Percentage of blacks	0.602***	0.322***	0.712***	0956**	0.591**	0.551**
	(0.194)	(0.112)	(0.224)	(0.485)	(0.285)	(0.265)
Growth rate of per capita GDP	-0.111	-0.058	-0.075	-0.152	-0.224*	-0.01
	(0.086)	(0.049)	(0.104)	(0.281)	(0.121)	(0.008)
Unemployment rate	0.012***	0.009***	0.015***	0.043***	0.018***	0.025***
	(0.003)	(0.002)	(0.004)	(0.011)	(0.005)	(0.005)
Percentage of college-educated	-0.632***	-0.412***	-0.765***	-0.755	-0.521**	-0.721**
	(0.153)	(0.088)	(0.191)	(0.457)	(0.198)	(0.18)
MSA dummies	Yes	Yes	Yes	Yes	Yes	Yes
	II. 2SLS first stage estimates					
(Imports from China to OTH)/worker	0.792***					
	(0.079)					
First-stage F-statistics	25					
Hausman test (t-statistics)	4.55	5.1	4.22	3.85	4.35	3.55
R-squared	0.14	0.135	0.185	0.168	0.175	0.183

Table 4: The Impacts of China Imports on Income Inequality: Subgroup I

The table shows the impacts of imports from China on income inequality for different age groups and education groups. Column (1) to (5) provide 2SLS second stage estimates while column (6) to (10) present OLS regression results. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Gini coefficient, (3) the natural logarithm of the Theil index, (4) the natural logarithm of the ratio of the 90th and 10th percentiles, and (5) the natural logarithm of the ratio of the 75th and 25th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. All regressions include MSA dummies and the MSA level controls. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25
	Panel I: By Age group									
	Age 16-34									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.011***	0.006***	0.012***	0.031***	0.013**	0.009	0.005	0.012	0.035***	0.012*
	(0.003)	(0.002)	(0.003)	(0.008)	(0.005)	(0.006)	(0.004)	(0.08)	(0.025)	(0.0066)
R-squared	0.14	0.135	0.185	0.168	0.175	0.132	0.12	0.176	0.17	0.182
	Age 35-49									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.012***	0.007***	0.0115***	0.0294***	0.010**	0.010	0.007*	0.01	0.025	0.008
	(0.003)	(0.002)	(0.003)	(0.008)	(0.005)	(0.006)	(0.004)	(0.007)	(0.02)	(0.03)
R-squared	0.13	0.14	0.192	0.172	0.16	0.135	0.13	0.16	0.17	0.2
	Age 50-64									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.014***	0.006***	0.013***	0.029***	0.011**	0.01	0.005	0.01	0.02	0.008
	(0.003)	(0.002)	(0.003)	(0.008)	(0.005)	(0.008)	(0.004)	(0.01)	(0.03)	(0.009)
R-squared	0.15	0.11	0.23	0.20	0.151	0.12	0.12	0.25	0.21	0.153
	Panel II: By education level									
	College									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.010*	0.003	0.013**	0.029	0.009	0.009	0.002	0.01	0.025	0.008
	(0.005)	(0.002)	(0.006)	(0.015)	(0.005)	(0.006)	(0.004)	(0.006)	(0.015)	(0.0056)
R-squared	0.12	0.12	0.14	0.19	0.15	0.156	0.18	0.13	0.18	0.13
	Non-College									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.02***	0.008***	0.015***	0.032***	0.02***	0.01	0.005	0.01	0.022	0.01
	(0.003)	(0.002)	(0.003)	(0.008)	(0.005)	(0.003)	(0.004)	(0.006)	(0.05)	(0.008)
R-squared	0.18	0.14	0.25	0.21	0.151	0.19	0.15	0.3	0.25	0.18

Table 5: The Impacts of China Imports on Income Inequality: Subgroup II

The table shows the impacts of imports from China on income inequality for different manufacturing sectors and sex groups. Column (1) to (5) provide 2SLS second stage estimates while column (6) to (10) present OLS regression results. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Gini coefficient, (3) the natural logarithm of the Theil index, (4) the natural logarithm of the ratio of the 90th and 10th percentiles, and (5) the natural logarithm of the ratio of the 75th and 25th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. All regressions include MSA dummies and the MSA level controls. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25
	Panel I: By Manufacturing Sector									
	Manufacturing Sector									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.013***	0.01***	0.02***	0.04***	0.02**	0.007	0.004	0.008	0.04***	0.013*
	(0.004)	(0.003)	(0.005)	(0.01)	(0.01)	(0.006)	(0.004)	(0.08)	(0.025)	(0.0066)
R-squared	0.142	0.136	0.19	0.17	0.12	0.13	0.11	0.16	0.15	0.19
	Non-Manufacturing Sector									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.008	0.002	0.004	0.016**	-0.01*	0.01	0.004	0.015	0.02	0.008
	(0.007)	(0.002)	(0.003)	(0.008)	(0.005)	(0.007)	(0.004)	(0.01)	(0.02)	(0.03)
R-squared	0.133	0.142	0.185	0.166	0.13	0.135	0.13	0.16	0.17	0.2
	Panel II: By Sex									
	Male									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.015***	0.02***	0.021***	0.042***	0.025**	0.008	0.003	0.007	0.03	0.01
	(0.003)	(0.002)	(0.006)	(0.01)	(0.01)	(0.006)	(0.004)	(0.006)	(0.015)	(0.02)
R-squared	0.14	0.152	0.185	0.16	0.13	0.156	0.18	0.13	0.18	0.13
	Female									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.014***	0.022***	0.025***	0.04***	0.03**	0.009**	0.004	0.005	0.01	0.012
	(0.004)	(0.003)	(0.004)	(0.01)	(0.01)	(0.004)	(0.004)	(0.006)	(0.05)	(0.008)
R-squared	0.15	0.133	0.22	0.19	0.14	0.19	0.15	0.3	0.25	0.18

Another concern for the validity of the analysis the time period that I examine: the increased imports from China could be a symptom of this increase of income inequality rather than a cause, namely the increase of income inequality coincides with the increased imports from China. For instance, the adoption of new technology creates or upgrades industries from labor-intensive to knowledge-intensive. The upgrading process amplifies the education premium and increases the income inequality. Meanwhile, for those firms in traditional labor-intensive industries, they would offshore jobs from the U.S. to China where labor costs is much cheaper. This automatically increases imports from China. Thus technological change is the drive behind both the increase of income inequality and imports from developing countries like China.

In order to relieve the concern and ensure that the results capture the period-specific effects of exposure to Chinese trade instead of some long-run common factors which drive both the rise of China import and increasing of income inequality, I conduct a falsification exercise by regressing *past* changes in the income inequality variables on *future* changes in import exposure. More specifically, I regress the change of income inequality variables 1993-2000 on the changes in import exposure 2000-2007. One would expect the positive association between the import exposure and income inequality should be absent in the falsification regression. In Table 6, the first five columns test the prediction using the quantile of MSAs with highest ratio of trade exposure in the 2000s relative to exposure in the 1990s, i.e. the local labor markets where exposure accelerated the most across the two time periods. And indeed, one can find that the re-distributive effect of import exposure is only significant for current period exposure and there is a week positive relationship between the income inequality and future exposure. The regression for the full sample of MSA in column (6) to (10) also confirm the same pattern.

Table 6: The Impacts of China Imports 2000-2007 on Income Inequality 1993-2000

The table shows the impacts of imports from China 2000-2007 on income inequality 1993-2000. Regressions in panel I include the quartile of MSAs with the largest ratio of import exposure 2000-2007 vs import exposure 1993-2000. The variable "future period imports" in panel B refers to a MSA's import exposure during the period 2000-2007. Column (1) to (10) provide 2SLS second stage estimates. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Theil index, (3) the natural logarithm of the ratio of the 90th and 10th percentiles, (4) the natural logarithm of the ratio of the 75th and 25th percentiles and (5) the natural logarithm of the ratio of the 90th and 50th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. All regressions include MSA dummies and the MSA level controls. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	I. MSAs w/strong growth of import exposure 2000-2007 versus 1993-2000					II.All MSA				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Logistic Gini	Log Theil	Log 90/10	Log 75/25	Log 90/50	Logistic Gini	Log Theil	Log 90/10	Log 75/25	Log 90/50
	2SLS second stage estimates									
Panel A. Current period exposure (1993-2000)										
(Imports from China to US)/worker	0.005***	0.02**	0.03*	0.008**	0.005	0.009***	0.025***	0.04***	0.01**	0.02***
	(0.002)	(0.001)	(0.016)	(0.003)	(0.008)	(0.004)	(0.003)	(0.012)	(0.005)	(0.007)
R-squared	0.18	0.23	0.25	0.3	0.21	0.146	0.24	0.188	0.29	0.21
Panel B. Future period exposure (2000-2007)										
(Imports from China to US)/worker	0.008	0.023	0.04	0.006	0.003	-0.01	0.025	0.03*	0.01	0.009
	(0.006)	(0.05)	(0.05)	(0.005)	(0.008)	(0.004)	(0.03)	(0.018)	(0.02)	(0.01)
R-squared	0.13	0.15	0.24	0.17	0.25	0.18	0.3	0.2	0.185	0.234

One of the motivations for my study to focus on Chinese imports is the observation that China acounts a large portion of the dramatic increase in US imports from low-income countries (Figure 1). One way to rule out the plausible argument that much of China's recent trade expansion has been driven by labor demand shocks in the United States is to consider Chinese imports alongside those of other countries. Table 7 presents the results, it compares the impact of growing exposure to Chinese imports to the effect of exposure to imports from other source countries. Panel I repeat the baseline estimates from Table 2 (OLS) and Table 3 (2SLS). Panel II shows the effect of imports from all low-income countries (China included) which is almost the same as the imports effect from China, it suggests that imports from other low-income countries may have a similar impact on US manufacturing as Chinese imports. Panel III and IV examines the impact on income inequality of imports from Mexico and Central America. One can observe again that by adding imports from Mexico and Central America to those of China does not affect the 2SLS estimates to China's imports alone, which indicates the idea that trade with China is the driving force behind supply-driven US imports from lower wage countries. In Panel IV which excludes China imports from Panel III, the estimates coefficients show a weakly positive relationship between trade exposure Mexico and Central America and income inequality. It lends strong supports that it was the dramatic increase of Chinese imports that is behind the increase of income inequality of US household. The final Panel estimates in Table 7, analyzing the impact of all other middle-income and high-income-country imports on US manufacturing, find small and inconsistently signed effects.

As mentioned in Autor, Dorn and Hanson (2013) is the measurement issue that imports from China contains both final goods purchased by US consumers and intermediate inputs purchased by US firms. Following Autor, Dorn and Hanson (2013), I subtract China imports of

Table 7: The Impacts of Imports on Income Inequality From Different Exporting Sources

The table shows the impacts of imports from different exporting sources on income inequality. Column (1) to (5) provide 2SLS second stage estimates while column (6) to (10) present OLS regression results. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Theil index, (3) the natural logarithm of the ratio of the 90th and 10th percentiles, (4) the natural logarithm of the ratio of the 75th and 25th percentiles and (5) the natural logarithm of the ratio of the 90th and 50th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. All regressions include MSA dummies and the MSA level controls. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25
	Panel I: China									
	2SLS second stage estimates					OLS				
(Imports from China to US)/worker	0.015***	0.008***	0.025***	0.038***	0.01**	0.009	0.005	0.01*	0.025*	0.008
	(0.006)	(0.003)	(0.006)	(0.015)	(0.005)	(0.006)	(0.003)	(0.06)	(0.015)	(0.005)
R-squared	0.14	0.135	0.185	0.168	0.175	0.133	0.132	0.17	0.354	0.285
	Panel II: China + other low income countries									
	2SLS second stage estimates					OLS				
(Imports from China										
+ other low income countries to US)/worker	0.013***	0.0075***	0.012***	0.029***	0.011**	0.012*	0.008**	0.02	0.033	0.06
	(0.003)	(0.002)	(0.003)	(0.008)	(0.005)	(0.006)	(0.004)	(0.006)	(0.03)	(0.05)
R-squared	0.135	0.142	0.191	0.175	0.165	0.137	0.131	0.162	0.175	0.22
	Panel III: China + Mexico/Cafta									
	2SLS second stage estimates					OLS				
(Imports from China										
+ Mexico/Cafta exporters to US)/worker	0.018***	0.009***	0.03***	0.044***	0.015***	0.008	0.01	-0.02	-0.05	0.01
	(0.004)	(0.003)	(0.004)	(0.008)	(0.005)	(0.01)	(0.05)	(0.05)	(0.05)	(0.02)
R-squared	0.154	0.113	0.235	0.2	0.155	0.12	0.12	0.25	0.21	0.153
	Panel IV: Mexico/Cafta									
	2SLS second stage estimates					OLS				
(Imports from Mexico/Cafta to US)/worker	0.006	0.001	0.009*	0.022	0.007	0.009	0.002	0.01	0.025	0.008
	(0.005)	(0.002)	(0.005)	(0.02)	(0.005)	(0.006)	(0.004)	(0.006)	(0.015)	(0.0056)
R-squared	0.121	0.08	0.11	0.2	0.3	0.156	0.18	0.13	0.18	0.13
	Panel V: all other exporters									
	2SLS second stage estimates					OLS				
(Imports from all other exporters to US)/worker	0.001	0.002	0.004	0.007	0.006	-0.002	-0.004	-0.01*	-0.022	0.009
	(0.003)	(0.002)	(0.003)	(0.008)	(0.005)	(0.003)	(0.004)	(0.005)	(0.05)	(0.008)
R-squared	0.11	0.15	0.21	0.22	0.15	0.12	0.16	0.33	0.25	0.19

Table 8: The Impacts of China Imports on Income Inequality: Other Measure of Import Competition

The table shows the impacts of imports from China on income inequality. The measures of income inequality are: (1) the logistic transformation of the Gini coefficient, (2) the natural logarithm of the Gini coefficient, (3) the natural logarithm of the Theil index, (4) the natural logarithm of the ratio of the 90th and 10th percentiles, (5) the natural logarithm of the ratio of the 75th and 25th percentiles ,and (6) the natural logarithm of the ratio of the 90th and 50th percentiles. Income inequality data are from the Current Population Survey (CPS). We use total personal income and sampling weights in the Current Population Survey (CPS) to calculate each inequality measure for each MSA in each year. The sample period is 1993 to 2007. First stage estimates in panel II also include the control variables that are indicated in the corresponding columns of panel I. Standard errors clustered by MSA are reported in parentheses, where *** indicates significance at the 1% level, ** at the 5% level, and * at the 10% level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Logistic Gini	Log Gini	Log Theil	Log 90/10	Log 75/25	Log 90/50
	2SLS second stage estimates					
(Imports from China to US net of intermediate inputs)/worker	0.008**	0.005***	0.022***	0.03**	0.008*	0.01*
	(0.004)	(0.001)	(0.003)	(0.012)	(0.004)	(0.008)
Percentage of blacks	0.61***	0.42***	0.812***	0.856**	0.791**	0.611**
	(0.19)	(0.12)	(0.242)	(0.425)	(0.235)	(0.245)
Growth rate of per capita GDP	-0.112	-0.06	-0.08	-0.155	-0.28*	-0.009
	(0.09)	(0.045)	(0.12)	(0.34)	(0.15)	(0.008)
Unemployment rate	0.012***	0.011***	0.02***	0.045**	0.021***	0.023***
	(0.0034)	(0.003)	(0.005)	(0.021)	(0.006)	(0.006)
Percentage of college-educated	-0.635***	-0.44***	-0.745***	-0.65	-0.531**	-0.745***
	(0.155)	(0.099)	(0.161)	(0.55)	(0.21)	(0.22)
MSA dummies	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.145	0.132	0.194	0.178	0.182	0.23

intermediate inputs per worker from total China imports per worker to mitigate the potential concern that the US firms' demand for China intermediate inputs might drive China's variety of inputs to US producers. Table 8 presents the results and one can see that while the magnitudes of coefficients have reduced (e.g. 0.008 column (1) vs 0.015 in column (1) of Table 3) but the estimates are still positively significant.

One important concern is the increase in inequality may be due to the migration flows. For instance, one might argue that higher exposure to Chinese trade might cause unskilled workers flow into states with lower exposure, which might change the income inequality among MSAs. Due the lack of data on the migration flows at MSA level, I can not empirically test the regional population shifts in response to trade shocks. One of the empirical analyses one can conduct to test the effect of migration flow is to recalculate the measures of inequality dropping from the sample all individuals who moved to the MSA during the previous year. However, this data is only available at a higher level at states or commuting zones.

2.6 Conclusion and Policy Implication

The starting point of this paper was two observations: firstly, the value of annual US goods imports from China increased by a staggering 1,156 percent from 1991 to 2007, whereas US exports to China grew by much less. And meanwhile, it is still a consensus (upon until recent) that trade has only modest effects on inequality rests on relatively old data. Understanding whether this increase has been a force for greater inequality in the United States and other developed countries requires more rigorous econometric design and update-to-date dataset.

In this paper, I re-consider the question by using the income inequality measures constructed from Current Population Survey (CPS) data and analyzing the effect of rising Chinese import competition between

1993 and 2007 on US local labor markets. Using a newly constructed measure of import exposure enables me to take advantage of the cross-market variation stemming from initial differences in industry specialization at MSA level. I instrument the import exposure using changes in Chinese imports by other high-income countries to overcome the potential endogeneity problems. I find that areas which are more exposed to China imports competition have larger changes in income inequality. In my main specification, a $1,000 exogenous decadal rise in a MSA's import exposure per worker leads to a 1.5% increase in the logistic Gini.

By revisiting the relationship between change in income inequality and trade, the contribution of this paper is at least two folds. Firstly, I believe that this evidence calls into question the consensus of the early 2000s, and reinforce that, after the early Bretton Woods era aberration, the distributional consequences of trade are effective. More importantly, it shows that the free trade among countries could have both benefits and costs. And the costs, especially trade adjustments costs, have only been recognized in theory and was not empirically evident. Thus a better understanding when and where trade is costly, and how and why it may be beneficial, are key items on the research agenda for trade and labor economists.

Theory suggests that trade with China yields aggregate gains for the US economy. My study provides more direct evidence on the distributional consequences of trade and the medium-run efficiency losses associated with adjustment to trade shocks. The results indicate that areas which are more exposed to China imports competition have bigger changes in income inequality. The consequences of China trade for US household income inequality may contribute to public ambivalence toward globalization and specific anxiety about increasing trade with China.

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Data Appendix

Measuring the industry structure of local labor markets

I derive the potential exposure of MSA to import competition from detailed information on local industry employment structure, which is taken from the County Business Patterns (CBP) data. CBP provides annual data series that contains information on employment, firm size distribution, and payroll by industry on country, MSA and state level. It contain all U.S. employment except self-employed individuals, employees of private households, railroad employees, agricultural production employees, and most government employees. CBP data is extracted by the U.S. Census Bureau.[1] The CBP data does not have information on individual employers, and information on employment by MSA and industry is thus sometimes reported as an interval instead of an exact number. Additional, some establishments are not identified at the most disaggregate level of the industry classification. Following Autor, Dorn and Hanson (2013), I calculated the imputed employment by MSA by 4-digit SIC code for years in 1980 and 1990 using the following steps:

[1]The data is available for downloading at `http://www.census.gov/econ/cbp/index.html`

- (1) Narrowing the range of possible employment values in cells with bracketed employment counts using the minimum and maximum employment values that are consistent with a cell's firm size distribution, and with the employment count of the corresponding aggregate industry.

- (2) Constructing a sample with all non-empty MSA-level 4-digit industry cells, and regress the employment in there cells on the number of firms in each of the 9 or 13 establishment size classes. The starting value of employment for cells with bracketed employment counts is the midpoint of the bracket. The coefficients of the regression yield an estimate for the the typical firm size within each firm size bracket. I replace employment counts in cells with bracketed values with the predicted values from the regression, and repeat the estimation and imputation until the coefficients of the establishment size variables converge.

- (3) Employing the establishment size information in 4-digit and corresponding 3-digit industries, and the coefficients from the preceding regression analysis to compute the employment in firms that are identified only by a 3-digit industry code in the data, and repeating the same step for higher levels of industry aggregation.

- (4) Proportionally adjusting estimated employment in 4-digit industries and in firms that lack 4-digit code so that they sum up to the employment of the corresponding 3-digit code. Repeat this step for higher levels of industry aggregation.

- (5) Assigning employment of firms that are only identified at the 2-digit industry level to 3-digit industries, proportional to observed 3-digit industry employment in the respective MSA. Repeat this step for assigning 3-digit employment to 4-digit industries.

Following the same procedure as outlined above for CBP 1990 and 1980, I calculate the imputed local labor structure for CBP 2000. The only difference is that CBP 2000 reports employment by MSA and industry for 6-digit NAICS codes and the distribution of firm sizes ranging from 9 to 13 size classes. For the concordance between NAICS and SIC codes, I employ a weighted crosswalk based on the Census "bridge" file available on the `Census.gov`. It reports the number of employees and

firms in the 1997 Economic Census for each existing overlap between NAICS and SIC industry codes.

Appendix Table 1
Different Measures of Income Inequality

Measures	Mathematical Expression	Interpretation	Advantages	Disadvantages
Gini coefficient	$1-2\int L(x)dx$, where $L()$ is the Lorenz curve showing the relation between the percentage of income recipients and the percentage of income they earn.	The Gini coefficient is equal to 0 in the case of perfect equality when exactly s percent of total income is held by bottom s individuals ($s=1,...,100$). The Gini coefficient is equal to 1 if all the income is held by one individual.	1. Very intuitive and widely used. 2. Makes use of all information about the distribution.	1. Sensitive to changes in the middle of the distribution. 2. Not easily decomposable to between- and within-group inequality
Theil index	$\frac{\sum_i \frac{y_i}{\mu}\ln\frac{y_i}{\mu}}{n}$, where i indexes individuals ($i=1,...,n$), y is personal income, and μ is the mean value of y. The first term inside the sum is individual's share of total income and the second term is that individual's income relative to the mean.	If all individuals have the same (i.e., mean) income, then the Theil index is 0. If one individual has all the income, then the index is $\ln(n)$.	Easily decomposable to between- and within-group inequality.	Hard to interpret.
Log(75/25)	$\ln(y75)$ $\ln(y25)$, where $y75$ and $y25$ are the 75th and the 25th percentiles of personal income distribution (y), respectively.	The ratio is equal to 0 if the 75th and the 25th percentiles of the distribution are equal. There is no upper bound to the ratio.	1. Intuitive measure of the percentage difference between the third and the first quartiles of a distribution. 2. Robust to extreme values.	Does not measure the entire distribution.
Log(90/10)	$\ln(y90)$ $\ln(y10)$, where $y75$ and $y25$ are the 90th and the 10th percentiles of personal income distribution (y), respectively.	The ratio is equal to 0 if the 90th and the 10th percentiles of the distribution are equal. There is no upper bound to the ratio.	1. Intuitive measure of the percentage difference between the top and the bottom deciles of a distribution. 2. Robust to extreme values	Does not measure the entire distribution.

Appendix Table 2

Imports from China and Other Developed Economics 1991 - 2007 (in Billion of 2007$) and their Correlations with U.S.-China Imports

	United States	Japan	Germany	Spain	Australia
Δ Chinese Imports (Bil$)	303.8	108.1	64.3	23.2	21.5
# Industries with Import Growth	385	368	371	377	378
Correlation w/U.S.-China Imports	1.00	0.86	0.91	0.68	0.96
	8 Non-US Countries	Finland	Denmark	New Zealand	Switzerland
Δ Chinese Imports (Bil$)	234.7	5.7	4.7	3.8	3.3
# Industries with Import Growth	383	356	362	379	343
Correlation w/U.S.-China Imports	0.92	0.58	0.62	0.92	0.55

Figure 1: U.S. Imports from China, Canada and Mexico/CAFTA)

This figure shows the U.S. trade with China and other countries, 1991-2007. The trade volume is in billion 2007 US$. The vertical line is the year 2001 when China joined World Trade Organization (WTO). The three lines stand from the U.S. trade with China (circle), Canada (diamond) and Mexico and the Central American and Carribean countries covered by the CAFTA-DR (triangle).

Figure 2.1

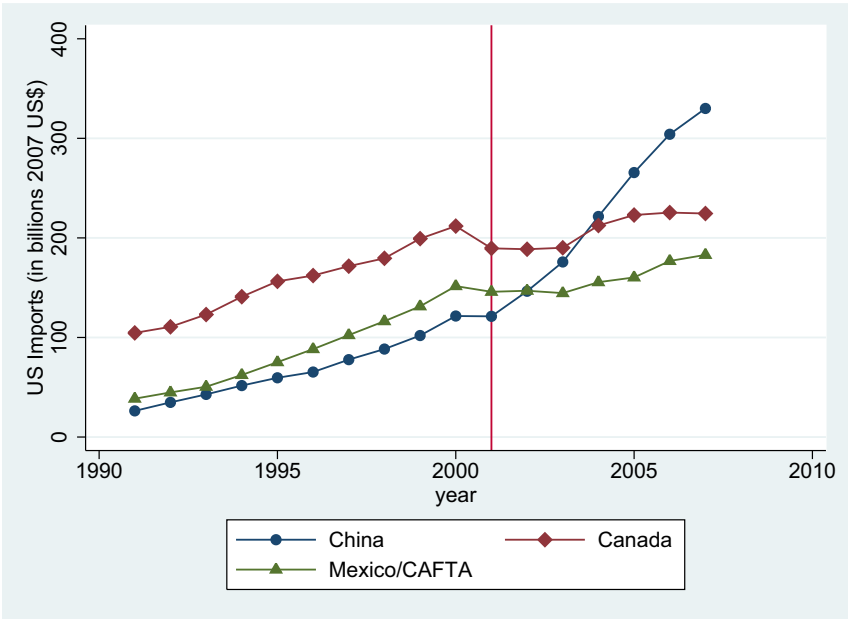


Figure 2: Relative Prices of Developing and Developed Country Manufactures, 1990 - 2014)

This figure shows the logarithm of this ratio between price indices of manufactured imports from developing and developed countries , normalized so that 1995=0. Data source: Krugman (2008) and Bureau of Labor Statistics.

Figure 2.2

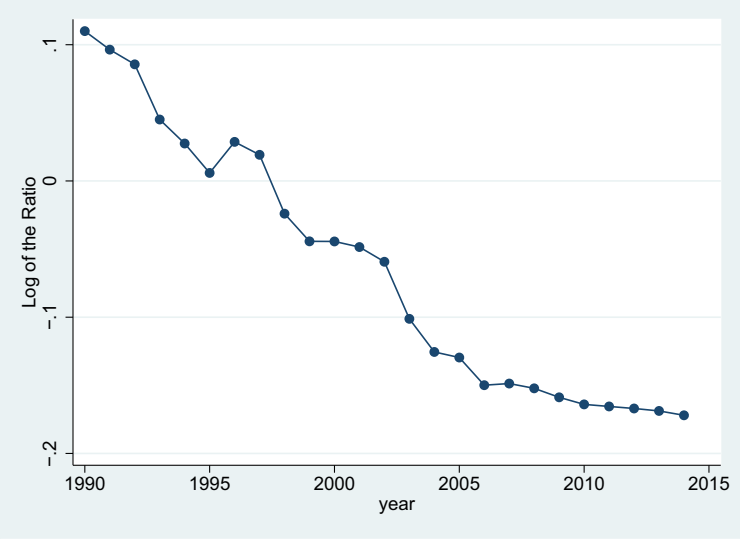
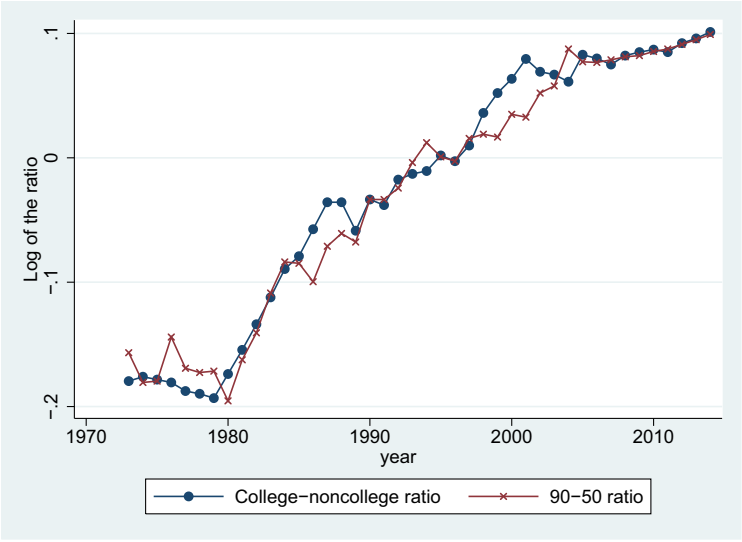


Figure 3: Hourly Wage Differentials for Males, 1973-2014

This figure shows shows two widely used indicators of wage differentials: the 90-50 ratio of hourly wages and the college-noncollege ratio. Data source: Krugman (2008) and Bureau of Labor Statistics.

Figure 2.3



3. Employee as Creditor: Evidence from Defined Pension Plans

In this paper, I show the role of pension plans in shaping the firms' labor market decision. By employing the loan covenants violation and consequently transferring of control rights to creditors, I examine the strategic use of pension underfunding by firms and the resultant wage cuts. I also find that the wage concession is less severe for firms from industry with bigger bargaining power. This study sheds light on how firms strategically renegotiate labor contracts to extract concessions from labor. The evidence suggests that credit contracts between debt-holders and shareholders have spillover effects on non-financial stakeholders.

3.1 Introduction

Pension arrangements in public and private sector employees have drawn increasing attention in recent years. During the financial crisis, retirement system has been overhauled to some extent especially in the U.S. On the one hand, there is significant changing from Pay-As-You-Go pensions towards defined contribution (DC) plans. On the other hand, defined benefit (DB) pension plans are usually being restructured and employed as a thread of 'pension dumping' in which during wage negotiation firms threaten to default on their pension obligations, thereby renegotiating a better terms (from the firms' perspective) of future retirement benefits to workers.[1]

The main purpose of this paper is to analyze the role of pension plan underfunding on firm's wage cut when the employees are seen as creditors of firms. The anecdotal evidence suggests that firms tend to negotiate down wages in times of financial distress. Following this point of view, firms with underfunding pension plan might have a bigger bargaining power when it comes to wage concession. Given the protection level of defined benefits of pension, it can be seen as unsecured debt of firm to employees. And firms in financial distress (default or close to bankruptcy) have strong motivation to strip off the pension from their balance sheet. In fact, firms can legally and easily dump the pensions especially when they are in the bankruptcy process (pension dumping). Therefore, afraid of losing their future income after retirement, employees are in the weak position in negotiating the wages and are willing to sacrifice their current income especially for those firms with deeply underfunded pensions plans.

The expropriation of employees by other stakeholders like manager and creditors shall be evident long beyond the firms' default or bankruptcy. Thus, the clear way to identify the conflicts are when firms

[1]Please see Hawthrone (2008) for a detailed description of pension dumping phenomenon.

are in financial distress or firms are controlled by other creditors. In this paper, I employ the identification by Chava and Roberts (2008) when the financial covenants are violated, the creditors (especially banks) take control the firms in question. It provides a quasi-natural experiment to analyze when secured creditors (primarily banks) gain control, how firms managers take the advantage of other stakeholders in order to cater big creditors. Previous studies show that when financial covenants are violated, firm reduces investment (Nini, Sufi and Smith (2009, 2012)), reduce innovation (Gu, Mao and Tian (2014)), and more likely to reduce workplace safety (Cohn and Wardlaw (2014)). The deficit of pension plan should be a sensitive subject for firms which are under control by creditors when it comes to firms' employment decision (wage cutting, scaling down the current pension plan or layoffs).

There is a rich and vast literature spanning several decades of studying defined-benefit (DB) corporate pension plans. The early contritions in this study mainly focuses on the implications of corporate taxes for the funding status of pension plans and for their asset allocation. Among them, for instance, Ippolito (1985) firstly points out that firms might intend to strategically scale up their pension plan deficit even in the presence of tax disadvantage. By doing so, they would be able to prevent the labor unions from holding-up the firm.

Even though the tax benefits for sponsors of DB pension plans are significant, the more recent literature has emphasized less on taxes as the main determinant of funding and pension plan asset allocation decisions. Instead, the focus has shifted to the guarantees provided by the government through the Pension Benefit Guarantee Corporation (PBGC), created in 1974 under the Employee Retirement Income Security Act (ERISA) to protect employee pensions from the risk of sponsor bankruptcy and default. Since then, a large body of studies has been conducted on the risk-shifting behavior that the provision of such insurance may induce. Sharpe (1976), Treynor (1977), and Harrison & Sharpe (1983)

are among the first papers which have attempted to quantify the value of this guarantees provided by the PBGC. Besides the policy implication itself, the introduction of the PBGC insurance has also proved beneficial in testing the well-established and understood concept in corporate finance: risk-shifting. For instance: Bodie et al. (1985, 1987), who document a negative relationship between underfunding and investment in bonds. Similarly, Petersen (1996) estimates a positive relationship between underfunding and investment in bonds. A more recent paper using variation in the degree of pension coverage provided by the PBGC, Benmelech, Bergman and Enriquez (2012) provides direct evidence on that actual mechanism in which airlines use underfunded pensions and the threat to dump those pensions in such way to extract labor concessions during the bankruptcy process. There is also evidence from Petersen (1992) who argues that firms decide to terminate their over-funded pension plans in order to transfer wealth from employees to shareholders. My paper distinguishes from prior studies by providing another mechanism. More specifically, it might not only be shareholders and their proxies in firms, managers but also other creditors would expropriate the employees' benefits during financial distress. This paper complements Liang and Falato (2014)'s finding that it was the pension plans underfunding plays a central role in negotiations between the employees and firms when the firms in question is under financial distress.

There is an emerging literature on the interaction between corporate finance and labor economics which argue that debt can be used strategically to control wage demands and that unionization levels are correlated with leverage at industry level (Baldwin (1983), Bronars and Deere (1991), Perotti and Spier (1993)). Empirical evidence like Hanka (1998) shows that debt and wage, employment and pension funding have a negative correlation. Klasa, Maxwell, and Ortiz-Molina (2008) find that firms in more unionized industries hold less cash. Employing exogenous vairation in state-level labor laws, Matsa (2009) shows that once

states adopt legislation which reduces union bargaining power, firms with concentrated labor markets reduce debt relative to other similar firms in other states. My paper adds to this literature in that by analyzing wage concessions, I identify a special condition under which the firms can use their financial situation to extract surplus from labor.

My work is also broadly related to several recent studies that examine the impact of pension funding on corporate decisions. For instance, Rauh (2006) employs the nonlinear funding rules for DB pension plans and finds a causal negative effect of mandatory pension contributions on corporate investments. By showing that managers adjust the assumed rates of return on pension assets in managing earnings, Bergstresser, Desai, and Rauh (2006) argue that the pension plan funding status affects managers' investment decisions. And Franzoni (2009) finds that stock prices do react to the payment of mandatory pension contributions to a firm's DB pension plan.

3.2 Institutional Background

Defined Benefit (DB) pension plans and Defined Contribution (DC) pension plans are the main types of institutional pension arrangements, which consist of the second pillar of retirement savings complemented by government-provided or state-run pensions (first pillar) and private savings (third pillar). For defined contribution plans, the employer only has a legal obligation to make specific contemporaneous payments into the pension account of the employee. Therefore the beneficiaries, i.e. the employees, bear the investment risk. On the other hand, a defined benefit plan specifics the benefits the employees receive at retirement, and the employer bears the investment risk. The employer is legally required to make contributions to the pension plan so that the assets are sufficient to meet the pension obligations. The analysis in this paper

focuses on DB plans, since the obligation of the employer is limited to periodic pension contributions in the case of DC plans. DB plans loss their popularity in the last decade to DC plans, due to the diminishing tax advantages, increased costs and competitive pressures given the few companies tend to adopt new DB plans. By saying that, DB plans however will not disappear completely in many cases for a long time into the future.

Pension plans sponsors are required by law to make mandatory contributions to their underfunded pensions plans. These mandatory pension are regulated under the rules established by the Internal Revenue Code and several pension protection acts that went through the changes over time, such as the Employment Retirement Income Security Act (ERISA) of 1974, the Pension Protection Act of 1987, the Retirement Protection Act of 1994, and the Pension Protection Act of 2006.

The Pension Benefit Guaranty Corporation (PBGC) - a federal corporation which protests the pensions of nearly 44 million American workers and serves as a statutory trustee of terminated pension plans when a sponsoring firm has insufficient assets to pay the benefits that participants are owed. If a firm filed for bankruptcy has insufficient pension assets to cover pension liabilities, the PBGC takes over the pension plan and guarantees that plan recipients receive their annual pensions up to a maximum number which is usually by law ($60,136 in 2015 for employees retiring at age 65, see Table A1 for a full detailed of maximum retiring guarantees set by PBGC.). Firms pay for this insurance with premiums that up to the number of covered employees and financing situation of the plans, but not the amount of risk taken with financial assets. During restructuring process after bankruptcy, the PBGC receives all the firm's pension assets. While most of defined benefits pensions in the U.S. are insured by the PBGC, this coverage is limited. The maximum annual gurantee is determined by employee age and was $60,136

Table 1: Variable Definitions

Vairable	Definition
Loan Covenant Violations:	
Bind	An indicator which equals to one if either net worth or current ratio fall below their respective loan covenant thresholds in any given firm-year.
Employment Outcome Variable:	
Employees	Number of employees for each firm.
% Change (wages/employees)	Percent change of average wage for each firm.
Firm Characteristics:	
Cash Flow	The ratio of income before extraordinary items plus depreciation and amortization over the net property , plant and equipment at the beginning of the fiscal year to total asset.
ROA	The ratio of operating income after depreciation over lagged total assets.
Assets	The book value of total assets. The natural logarithm of the book value of assets is used in regressions.
Investment	Capital expenditures over net property, plant and equipment at the beginning of the fiscal year.
Leverage	Book value of debt divided by book value of total assets.
Pension Underfunding Variabls:	
Underfunding dummy	An indicator which equals one if pension plan liability is larger than the pension plan asset in the current fiscal year.
Underfunding $>$ 10% of plan assets	An indicator which equals one if pension plan liability is larger than the pension plan asset in the current fiscal year.
Underfunding $>$ 15% of plan assets	An indicator which equals one if pension plan liability is larger than the pension plan asset at 15% in the current fiscal year.
Underfunding $>$ 20% of plan assets	An indicator which equals one if pension plan liability is larger than the pension plan asset at 20% in the current fiscal year.
Underfunding $>$ 25% of plan assets	An indicator which equals one if pension plan liability is larger than the pension plan asset at 25% in the current fiscal year.
Industry Characteristics:	
Unionization Membership	Share of employees in any given industry-year that are members of a union.
Unionization Coverage	share of employees in any given industry-year that are covered by a collective bargaining agreement.

for a 65 year-old employee in 2015[1]. Therefore, generally speaking, pension deficits are considered to be sponsoring firms' unsecured senior inside debt owed to employees, which are partly insured by the PBGC.

3.3 Data and Sample

My sample consists of all Compustat[1] firms incorporated in the United States that have relevant loan covenant information from Loan Pricing Corporation's (LPC) Dealscan database[2] for the period 1994 to 2010. Data from Compustat are merged with covenant violation information from Dealscan by matching company names and loan origination dates from Dealscan to company names and corresponding active dates in the CRSP historical header file. The basic unit of observation in Dealscan is a loan, also referred to as a facility or a tranche. Loans are often grouped together into deals or packages. Since information on covenants is limited prior to 1994, I focus my attention on the sample of loans starting from year 1994. A firm is in violation of a covenant if the value of its accounting variable breaches the covenant threshold - i.e., when the ratios below the corresponding threshold.[3] I focus on the net worth and

[1]Studies also indicate that pension deficits are far from fully guaranteed by the PBGC. For example, *Reuters* reports that "...the underfunded liability was estimated at $41 billion for GM, Chrysler and Ford Motor Co. at the end of 2008, the latest PBGC figures show. GM accounted for half the shortfall and only $4 billion of that gap would be insured if plans were terminated now, according to the PBGC. Chrysler had a $9 billion shortfall of which $2 billion would be covered." (*Reuters*, April 23, 2009)

[1]Compustat is a database of financial, statistical and market information on active and inactive global companies throughout the world. In this paper, I only use the public firms data from Compustat North America.

[2]The Dealscan database from Thomson Reuters contains historical information on the terms and conditions of deals in the global commercial loan market. It contains extensive data on detailed terms and conditions on over 200,000 loan transactions.

[3]The covenant violation data is provided by Michael Roberts in his personal website (http://finance.wharton.upenn.edu/ mrrobert/styled-9/styled-

current covenants for two reasons: firstly, these two types of covenants have the most frequent appearance in the Dealscan database. According to Chava and Roberts (2008) that covenants regarding the current ratio or net worth show up in 9294 loans with a combined face value of over a trillion dollars. Secondly and more importantly, the accounting measures used for these two covenants are standardized and consistent. In this paper, I do not discriminate between those two covenants, thus there is a violation if either of the two covenants is breached in any previous given firm-year.

The accounting items related to DB pension plans and financial statement data are drawn from Compustat Pension Annual dataset. Firms are defined to have DB plans both the fair value of pension plan assets (Compustat data item *pplao*) and the projected pension benefit obligation (Compustat data item *pbpro*) are available in Compustat[4]. The key variable of my interest, the pension underfunding, is defined as the difference between *pplao* and *pbpro*. Thus a positive (negative) pension funding status indicates that firm's DB pension plan is overfunded (underfunded). A negative pension funding status represents a true liability for the sponsoring company even if it does not appear on the balance sheet.

The number of employees and wage data is from Compustat as well. While Chava and Roberts (2008) primarily use quarterly data, I use annual data because firms do not report employment and wage at the quar-

11/index.html) with the together with the corporate ID (gvkey) and violation dates information.

[4]U.S. Internal Revenue Service (IRS) Form 5500 is another data source for a firm's pension-related variables. It provides detailed information regarding the plan's finances, participants, and administrators. Although the complete filings of Form 5500 are available for fiscal years 1990 through 2007 (http://crr.bc.edu/data/form-5500-annual-reports/), it is only the filings before 1998 provide the pension administrators (corporate) ID which enables me to merge with Compustat dataset. Thus using Form 5500 would significantly reduce my sample size.

terly frequency. I acknowledge that this data limitation does not perfectly fit my assessment when the covenant violation occurs more noisy, although Chava and Roberts document that their results also hold with annual data. I calculate wage average of all employees for each of firms and use the percent annual change in the average wages for each firms as my main dependent variable in the regression analysis.

Table 2 provides summary statistics for the incidence of loan covenant violations as well as medians, 25^{th} and 75^{th} of my main dependent variable and standard firm and industry characteristics in the resulting sample of 13077 firm-year observations that are bound by either a current ratio or a net worth covenant during the period 1994 to 2010. Overall, my sample is comparable to those used in previous studies (Chava and Roberts (2008), Dichev and Skinner (2002)). The frequency of firm-year observation which are categorized to be in violation is 20%, which is in line with the 15% frequency reported in Table 3 of Chava and Roberts (2008), I interpret the difference as the fact that my sample frequency is annual while theirs is based on quarter data and my longer sample period includes the financial crisis. Table 3 further compares the firm characteristics between firms with under-funded pension and non-underfunded pension. It is clear from the table that these two group are identical in terms of firm characters.

I further employ the representation of unionization dataset to analyze the cross-sectional effect of the wage concession through the channel of pension underfunding. I use the union membership and coverage database from CPS provided by Hirsch and Macpherson (2003) from the website `www.unionstats.com`. It is an Internet data resource providing private and public sector labor union membership and coverage[5] complied from the monthly household Current Population Survey (CPS) us-

[5]Coverage denotes the individual who are covered by a collective bargaining agreement while it is not necessary for him or her to be a labor union member. Thus the coverage and union membership are differentiated in the CPS survey questions.

Table 2: Summary Statistics

Variable	Mean	Median	SD	25^{th}	75^{th}	N
Loan Covenant Violations:						
Bind	0.21	0	0.12	0	0	13077
Employment Outcome Variable:						
Employees	15.9	3.27	40.8	0.77	12.2	13077
% Change (wages/employees)	0.17	0.13	0.11	0.077	0.25	13077
Firm Characteristics:						
Cash Flow	0.046	0.072	0.12	0.02	0.12	13077
ROA	0.096	0.11	0.11	0.06	0.16	13077
Assets	11750.9	848.2	78153.6	181.1	3834.2	13077
Investment	0.21	0.16	0.18	0.097	0.26	12666
Leverage	0.94	0.47	29.1	0.34	0.61	12080
Pension Underfunding Variabls:						
Underfunding dummy	0.54	1	0.5	0	1	13077
Underfunding $>$ 10% of plan assets	0.4	0	0.49	0	1	13077
Underfunding $>$ 15% of plan assets	0.35	0	0.48	0	1	13077
Underfunding $>$ 20% of plan assets	0.3	0	0.46	0	1	13077
Underfunding $>$ 25% of plan assets	0.26	0	0.44	0	1	13077
Industry Characteristics:						
Unionization Membership %	12.9	9.2	11.8	3.7	19	10460
Unionization Coverage %	13.9	10.2	12.2	4.3	20.4	10460

Table 3: Underfunded vs Non-underfunded Summary Statistics

	Underfunded		Non-underfunded		
Variable	Mean	N	Mean	N	Difference
Employees	15.52	7062	16.21	6015	-0.69
					(0.52)
% Change (wages/employees)	0.155	7062	0.182	6015	-0.027
					(0.031)
Cash Flow	0.044	7062	0.05	6015	-0.006
					(0.01)
ROA	0.11	7062	0.08	6015	0.03
					(0.03)
Assets	12,550.2	7062	11,880.3	6015	669.9
					(782.3)
Investment	0.18	7062	0.24	6015	-0.06
					(0.05)
Leverage	0.88	7062	1.02	6015	-0.14
					(0.12)

ing BLS methods. Due to the inconsistent industry codes among different waves of surveys and the differences from common industry codes in Compustat, I clean the data in the following steps: firstly, in order to have a consistent industry code for the union dataset, I transfer all the industry codes in the union database into 2002 survey of CPS industry codes (where it uses Census Industry Codes(CIC)). Secondly, using the names of industries, I manually merge the CIC codes with the corresponding North American Industry Classification System (NAICS), which is one of the industry identifications available in Compustat. The tricky part the process is that one CIC category might have several corresponds in NAICS, I start with finest level of NAICS (six digits), and then goes up one level (five digits) if the NAICS does not have the exact corresponds in CIC until a CIC code is matched to the NAICS in question. In the end, roughly 80% of the entire sample is able to match with unionization data.

I employ the UCLA-LoPucki Bankruptcy Research Database (BRD) to identify firms filed for Chapter-11 bankruptcy code. BRD is a data collection, data linking, and data dissemination project of the UCLA School of Law. It contains data on all of the more than one-thousand large public companies that have filed bankruptcy cases since October 1, 1979. Coverage includes cases filed under Chapter 7 and Chapter 11, whether filed by the debtors or creditors.

3.4 Empirical Strategy

To analyze the role of pension plan underfund in determining firms' labor market decision, I examine conflicts of interest creditors and an important class of non-financial stakeholders. Specifically, I assess the impact of loan covenant violations on employees. Why would covenant violations affect employees in the first place and how it would work

through the channel of pension underfunding? Loan covenants which are attached to performance indicators protect creditors by defining a transfer of control rights when a covenant is violated (see Aghion and Bolton (1992) and Dewatripont and Tirole (1994)). Such violations can be seen as a weak form of bankruptcy which gives creditors with the same rights as if the firms had default payments. Those rights usually include accelerating any outstanding principal payments or to entirely terminate any unused revolving loans. Managers of firms may decide to reduce operating costs by cutting jobs or renegotiate down wages after a covenant violation in order to avoid the consequences such as acceleration or termination of assessing to credit.[1]

Since employees with underfunded pension plans bear a higher cost when firms default, their outside options in the event of default is reduced. Therefore, in bargaining, management can employ the threat of 'pension dumping' to extract greater concession from labor. My identification then relies on a difference-in-difference specification, with two different lays of difference: (i) firm with covenants violations vs. firms without covenants violations, (ii) underfunded pension plans vs funded plans. I expect that while experiencing a covenants violation, firms can negotiate down the wages of their employees whose pensions are under-

[1] I provide an example of how employees may be affected by creditors' interventions. The following example is from the first quarter 10-Q filing of Interpharm Holdings in 2008: "Subsequently, on January 28, 2008, Wells Fargo informed the Company that it would consider providing the Company with credit availability on the condition that the Company (i) develops and implements a new operating plan focused on increasing the amount of eligible collateral and reducing costs and (ii) develop an alternative financing arrangement. Further, on February 5, 2008, the Company and Wells Fargo entered into the Forbearance Agreement [. . .] In connection with its negotiation of the Forbearance Agreement, the Company completed a restructuring of its operations on January 25, 2008 and submitted a new operating plan to Wells Fargo, which the Company believes will result in positive cash flow and net profits, and includes [. . .] reducing **payroll** and **headcount** by approximately 20%."

funded. More specifically, I am testing the following regression:

$$\Delta \frac{wages}{employees}_{i,t} = \beta_0 + \beta_1 Bind_{i,t-1} + \beta_2 underfunding_{i,t-1} + \beta_3 Bind_{i,t-1} \times underfunding_{i,t-1} + \beta_4 X_{i,t-1} + \eta_i + \lambda_t + \nu_{i,t} \tag{3.1}$$

where $\Delta \frac{wages}{employees}_{i,t}$ is the annual percent of change in the average wage of each firm. Subscripts indicate firm (i) and year (t); $Bind_{i,t-1}$ is a covenant violation dummy which equals to one if there is a violation in previous year, $underfunding_{i,t-1}$ is a dummy variable that equals to one for underfunded pension plans, $X_{i,t-1}$ is a vector of control variables measured at the fiscal year-end prior to the year in which wage is measured, η_i is a firm fixed effect, λ_t is a year fixed effect, and $\nu_{i,t}$ is a random error term assumed to be correlated within firm and potentially heteroskedastic. Controls include variables that have been previously employed in the loan covenants literature, for instance, firm size, profitability, and operating performance. β_3 is the parameter of interest, which represents the impact of a covenant violation on wage through the channel of underfunding pension plans.

Figure 1 plots the percentage annual changes in the average wage in event time leading to and after the year when a covenant violation occurs. The first quantile firms are the ones with the lowest underfunding status and the fourth quantile firms are the ones with the highest underfunded pension plans. The underfunding status is measured by the difference between pension plan liability and pension plan asset in the previous fiscal year. Two points are worth noticing from the graph. Firstly, it is clear from the figure that firms with in high and low underfunding status had a parallel trend in terms of percentage annual changes in the average wage before the covenant violation occurs, more precisely they are overlapping with each other in most cases. This suggests that the parallel trends hypothesis underlying our diff-in-diff framework is valid in the data. Secondly, average wage in the two groups increase

after the violation occurs but the effect is much larger for the firms in the fourth quantile. This figure provides compelling evidence that the covenant violation has had a large impact on firm's wage mainly due to its impact on underfunding situation by firms.

Figure 2 plots firms' financial characteristics in event time leading to and after the year when a covenant violation occurs. The first quantile firms are the ones with the lowest underfunding status and the fourth quantile firms are the ones with the highest underfunded pension plans. It is obvious from the figures that the patterns are identical for the two groups in terms of other firm's financial characteristics. More specific, cash flow has been decreasing but recovered quickly from the covenant violation. The covenant violation appear to have a relatively longer effect on debt issuing and investment: they have not had picked up five years after the violation occurs. Figure 2 has two implications: firstly it seems to suggest that covenants violations have had an effect on firm's performance (ROA), debt capacity (leverage), investment and cash holdings. And secondly, the violation does not impact the firm's performance heterogeneously through firm's pension underfunding situation.

3.5 Baseline Results

Table 4 reports the estimates from equation (1) in the entire sample. All the specifications include year and firm fixed effects. I firstly replicate the results of Chava and Roberts (2008) using investment as the dependent variable in my sample. In column (1) of Table 4, it clearly shows that investment is negatively significant at 5% level significance. As the columns (2) to (4) of Table 4 demonstrates, I find that the interaction term between covenant violation and underfunding dummy is consistently negative and statistically significant - firms which have a

Table 4: Loan Covenant Violations, Pension Underfunding and Wages I

This table presents the baseline DID specification with investment and percent change in average wages as the dependent variables. The firm accounting data is from Compustat. The dependent variable in column (1) is the investment which is defined as capital expenditures over net property, plant and equipment at the beginning of the fiscal year. The dependent variable from column (2) to (4) is the percent change in average wages. ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	(1)	(2)	(3)	(4)
	Investment	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %
Bind	-0.024**	-0.017	-0.015	-0.010
	(0.012)	(0.015)	(0.012)	(0.015)
Underfunding		0.001	-0.006**	-0.003
		(0.004)	(0.002)	(0.003)
Bind*Underfunding		**-0.110*** **	**-0.116*** **	**-0.115*** **
		(0.019)	**(0.016)**	**(0.019)**
Asset	0.016***	0.004	-0.005***	0.004
	(0.006)	(0.003)	(0.001)	(0.003)
Cash Flow	0.130***	-0.078***	-0.094***	-0.078***
	(0.035)	(0.022)	(0.017)	(0.022)
ROA	0.157***	-0.128***	-0.122***	-0.132***
	(0.056)	(0.028)	(0.019)	(0.028)
Firm Fixed Effects	Yes	Yes	No	Yes
Industry Fixed Effects	No	No	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	No
Observations	12,601	13,077	13,077	13,077
R-squared	0.587	0.470	0.240	0.465

Table 5: Loan Covenant Violations, Pension Underfunding and Wages II

This table presents the DID specification with percent change in average wages as the dependent variables. The firm accounting data is from Compustat. The dependent variable from column (1) to (4) is the percent change in average wages. Underfunding dummy is an indicator variable that equals one if underfunding is larger than 10% (column 1), 15% (column 2), 20% (column 3) and 25% (column 4). ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	Underfunding >			
	10%	**15%**	**20%**	**25%**
	(1)	(2)	(3)	(4)
	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %
Bind	-0.061***	-0.068***	-0.071***	-0.079***
	(0.015)	(0.015)	(0.014)	(0.015)
Underfunding	-0.002	-0.005	-0.004	-0.005
	(0.004)	(0.004)	(0.004)	(0.004)
Bind*Underfunding	**-0.068*** **	**-0.061*** **	**-0.060*** **	**-0.046** **
	(0.020)	**(0.021)**	**(0.021)**	**(0.022)**
Asset	0.004	0.004	0.004	0.004
	(0.003)	(0.003)	(0.003)	(0.003)
Cash Flow	-0.078***	-0.078***	-0.077***	-0.076***
	(0.023)	(0.023)	(0.023)	(0.023)
ROA	-0.133***	-0.133***	-0.133***	-0.134***
	(0.028)	(0.028)	(0.028)	(0.028)
Firm Fixed Effects	Yes	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No
Year Fixed Effects	Yes	Yes	Yes	Yes
Observations	13,077	13,077	13,077	13,077
R-squared	0.469	0.469	0.468	0.468

violation in covenants in previous years obtain wage concession from the employee whose pension plans that are underfunded.[1] My results hold after controlling for firm specific controls, firm fixed effects, year fixed effects and industry fixed effects. The estimates of β_3 are between -11% and -12%, suggesting that firms obtain wage concessions that are around 11 percent of annual wage when both covenants are violated and the employees' pension plan is underfunded.

The magnitude of the finding is larger than what Benmelech, Bergman and Enriquez (2012) finds. They estimate a 4 percent drop in annual average wage in airline industry. In their paper, they only consider the wage concessions when airlines have both a low cash flow and underfunded pension plan. A firm with a low cash flow is usually considered as a weak form of being financial constrained, while a covenants violation has a stronger effect and it usually induces the big creditor's intervenience. My result reinforces the role of big creditors in intervening firms' managerial and operating decisions. It well corresponds the conclusion in Liang and Falato (2014) where they find a substantial job cut of 19% per year after the bleach of covenants.

In Table 5 I refine the definition of underfunding dummy to capture higher level of underfunding. The difference-in-difference estimator is 7% when the pension plan is underfunded by at least 10%, while it is 6%, 6% and 4.6% when the pension plan is underfunded by at least 15%, 20% and 25%, respectively. It appears to suggest that the wage concession is not growing alongside with the level of pension plan underfunding. It also relieves the concerns that all effects come from some extreme underfunded firms.

[1]The coefficients of both covenants violations and underfunding are negative but not statistically different from zero in most of the specifications.

3.6 How does Labor Bargaining Power Affect the Wage Concession?

Labor unions are important determinants of wage increases and collective bargaining negotiations (Lewis (1986)). I expect the impact of covenant violation on wage concessions should be larger whenever employees are in a relatively weaker bargaining position with respect to creditors, since financial contracting theory suggests that employment cut are brought by the strengthening of creditor right relative to labor rights. In Benmelech, Bergman and Enriquez (2012), they only focus on a single industry (airline) which prohibit them to study the cross-sectional effect of labor unionization.

Table 6 examines variation by labor union coverage and membership representative as industry-level proxies of employee's bargaining power. In each year of the sample period, I rank firms based on the empirical distribution of these proxies. It has been well established that organized representation through unions can increase labor bargaining power (Clark (1984), Hirsch (2008)). I thus expect the impact of violations on wage concession when firms are underfunded should be strongest for firms in industries with the weakest bargaining power. Table 6 reports the results. Column (1) and (2) rank the sample using labor union membership and column (3) and (4) are ranked by labor union coverage. It seems to suggest that firms that are from a industry with the lower representative in labor unionization have the a bigger wage concession with underfunded pension plans.

Table 6: Loan Covenant Violations, Pension Underfunding and Wages: Labor Union Bargaining Power

This table reports the subsample regression results for firms that have high/low labor union representation. The firm accounting data is from Compustat. The dependent variable from column (1) to (4) is the percent change in average wages. A firm is identified as having a low (high) labor union representation if its industrial labor union representation in year t-1 is in the bottom (top) tercile among sample firms. Column (1) and (2) rank the sample using labor union membership and column (3) and (4) are ranked by labor union coverage. The labor unionization data is from Hirsch and Macpherson (2003). ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	(1)	(2)	(3)	(4)
	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %
	Membership		Coverage	
	Low	High	Low	High
Bind	-0.005	-0.037*	-0.009	-0.024
	(0.030)	(0.021)	(0.031)	(0.021)
Underfunding	0.001	-0.007	-0.003	-0.008
	(0.008)	(0.006)	(0.008)	(0.006)
Bind*Underfunding	-0.076**	0.028	-0.064*	0.017
	(0.035)	(0.026)	(0.036)	(0.026)
Asset	0.007	-0.001	0.005	-0.001
	(0.005)	(0.004)	(0.005)	(0.004)
Cash Flow	-0.138***	-0.086***	-0.153***	-0.084***
	(0.047)	(0.028)	(0.046)	(0.028)
ROA	-0.082	-0.003	-0.052	-0.002
	(0.058)	(0.037)	(0.057)	(0.037)
Firm Fixed Effects	Yes	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No
Year Fixed Effects	Yes	Yes	Yes	Yes
Observations	2,626	3,004	2,611	3,014
R-squared	0.621	0.591	0.609	0.587

3.7 Do Pension Plan Underfunding Still Matter For Firms in Chapter-11?

I move forward to analyze the effect of Chapter-11 on wage negotiation using the DD approach. I repeat the analysis in regression equation (1) but replace the loan covenant violation dummy with a bankruptcy dummy. Similar to our previous analysis, I estimate different specifications of the following baseline regression:

$$\Delta \frac{wages}{employees}_{i,t} = \beta_0 + \beta_1 Bankruptcy_{i,t-1} + \beta_2 underfunding_{i,t-1} + \beta_3 Bankruptcy_{i,t-1} \times underfunding_{i,t-1} + \beta_4 X_{i,t-1} + \eta_i + \lambda_t + \nu_{i,t}, \tag{3.2}$$

where $\Delta \frac{wages}{employees}_{i,t}$ is the annual relative change in the average wage of each firm. Subscripts indicate firm (i) and year (t); $Bankruptcy_{i,t-1}$ is a dummy variable which equals to one if a firm filed for a bankruptcy under Chapter-11 code in year $t-1$, $underfunding_{i,t-1}$ is a dummy variable that equals to one for underfunded pension plans, $X_{i,t-1}$ is a vector of control variables measured at the fiscal year-end prior to the year in which wage is measured, η_i is a firm fixed effect, λ_t is a year fixed effect, and $\nu_{i,t}$ is a random error term assumed to be correlated within firm and potentially heterokedastic. Controls include variables like firm size, profitability, and operating performance. β_3 is the parameter of interest, which represents the impact of a covenant violation on wage when firm is in the process of bankruptcy.

As mentioned before, the bleach of covenant can be seen as a weak form of a bankruptcy. From this perspective, firms in bankruptcy process would undertake a more serious restructure of the capital structure and thus creditors might as well employ the pension underfunding to negotiate down the wage of employees. On the other hand, current U.S. Bankruptcy Code is clearly in favor of maintaining the ongoing busi-

ness and workers' jobs rather than fulfilling firms' underfunded pension plan. In practice, bankruptcy courts rarely prioritize employees' claims on retiree benefits over other more senior creditors' claims or the possibility of the firm being restructured as a whole. The Horizon Natural Resources Company, an Illinois-based coal corporation can serve as an example.

> "...massive retiree benefit claims had to be dealt with if the company were to be sold as an entire reorganized unit. The retiree claims, however, were so daunting that no buyers wanted to touch the company. Finally, the bankruptcy court allowed the company to sell certain of its mines free and clear of the retiree benefit claims....that sales at least allowed some of the company's current workers to retain their jobs, albeit at reduced wage rates.[1]"

Therefore, the effect of bankruptcy on wage concession might not take effect from the channel of underfunded pension plans.

Table 7 reports the results from estimating the regression (2). I find that β_3 is negative but not statistically significant in all levels of underfunding. My result potentially indicates that firms tend to do most of the labor renegotiations before they file for Chapter-11 and they do not appear to strategically renegotiate wages during bankruptcy. My results tend to be favor of the hypothesis that the pension underfunding status is not the first order importance of channel in the determinant of wage concession during the firms' bankruptcy process.

[1] In re Horizon, 316 B.R. at 271, 283.

Table 7: Loan Covenant Violations, Bankruptcy and Wages

This table presents the DID specification with percent change in average wages as the dependent variables. The firm accounting data is from Compustat. The dependent variable from column (1) to (5) is the percent change in average wages. Underfunding dummy is an indicator variable that equals one if underfunding is larger than 0% (column 1), 10% (column 2), 15% (column 3), 20% (column 4) and 25% (column 5). Bankruptcy is a dummy variable which equals to one if the firm filed for bankruptcy in the previous year. ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	Underfunding >				
	0%	**10%**	**15%**	**20%**	**25%**
	(1)	(2)	(3)	(4)	(5)
	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %
Underfunding	-0.004	-0.006	-0.008*	-0.008*	-0.008**
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Bankruptcy * Underfunding	**-0.020**	**-0.022**	**-0.003**	**0.004**	**-0.008**
	(0.057)	**(0.041)**	**(0.047)**	**(0.050)**	**(0.053)**
Asset	0.006*	0.006*	0.006*	0.006*	0.006*
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Cash Flow	-0.070***	-0.070***	-0.070***	-0.070***	-0.069***
	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)
ROA	-0.124***	-0.124***	-0.124***	-0.124***	-0.124***
	(0.029)	(0.029)	(0.029)	(0.029)	(0.029)
Firm Fixed Effects	Yes	Yes	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No	No
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Observations	13,077	13,077	13,077	13,077	13,077
R-squared	0.453	0.453	0.454	0.454	0.454

3.8 Robustness Tests

3.8.1 Potentially Confounding Factors

To improve the validity, as well as the precision, of the estimates, I controlled for some major firms' characteristics which include firms' asset, cash flow and return-to-asset ration. Firm's asset proxies for the firm sizes, larger firms might have less variation in performance during the financial stress and thus might have less volatile in change of wage concession; cash flow is a direct measure for firms' financial stress, firms' with low cash flow usually indicates the firms' poor financial condition; and return-to-asset provides a direct measure for firm's profitability.

In Table 8, I present various specifications to compare with the baseline regression (Column (1)). In Column (2), I use the firm's asset with sale to proxy for their size. I use the KZ-index (Kaplan-Zingales Index) to measure firm's relative reliance on external financing. Companies with a higher KZ-Index scores are more likely to experience difficulties when financial conditions tighten since they may have difficulty financing their ongoing operations.[1] Finally, Tobin's Q is employed to measure the firm's performance as a replacement for ROA.

When I replace all the original proxies with the alternative ones, the interaction term of covenant violation dummy and pension underfund-

[1]The KZ-Index (Kaplan-Zingales Index) is based on the following five-factor model :
$KZ\ Index = -1.001909 \times Cash\ Flows/K + 0.2826389 \times Q + 3.139193 \times Debt/Total\ Capital + -39.3678 \times Dividends/K + -1.314759 \times Cash/K.$
$Cash\ Flows = Income\ Before\ Extraordinary\ Items_t + Total\ Depreciation\ and\ Amortization_t.$
$K = PP\&E_{t-1}.$
$Q = (Market\ Capitalization_t + Total\ Shareholder's\ Equity_t - Book\ Value\ of\ Common\ Equity_t - Deferred\ Tax\ Assets_t)/Total\ Shareholder's\ Equity_t.$
$Debt = Total\ Long\ Term\ Debt_t + Notes\ Payable_t + Current\ Portion\ of\ Long\ Term\ Debt_t.$
$Dividends = Total\ Cash\ Dividends\ Paid_t.$
$Cash = Cash\ and\ Short - Term\ Investments_t.$

Table 8: Stability Tests

This table presents the DID specification with percent change in average wages as the dependent variables. The firm accounting data is from Compustat. The dependent variable from column (1) to (5) is the percent change in average wages. Sale is the natural logarithm of firms' annual revenue. KZ-index (Kaplan-Zingales Index) is a measure of firm's relative reliance on external financing. Tobin's Q is firms' market value over book value. ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	(1)	(2)	(3)	(4)	(5)
	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %
Bind	-0.024	-0.031*	-0.033	-0.045	-0.033
	(0.017)	(0.016)	(0.022)	(0.035)	(0.032)
Underfunding	0.004	-0.003	0.003	0.002	-0.002
	(0.004)	(0.003)	(0.004)	(0.005)	(0.003)
Bind*Underfunding	-0.103***	-0.115***	-0.102***	-0.115***	-0.122***
	(0.021)	(0.038)	(0.033)	(0.038)	(0.032)
Asset	0.003		0.004	0.003	
	(0.003)		(0.003)	(0.003)	
Cash Flow	-0.059**			-0.066***	-0.063***
	(0.024)			(0.021)	(0.019)
ROA	-0.147***				-0.155***
	(0.031)				(0.029)
Sale		0.026			0.011
		(0.016)			(0.023)
KZ-index		0.063**	0.053***		
		(0.032)	(0.019)		
Tobin's Q		-0.225***	-0.221***	-0.222***	
		(0.071)	(0.069)	(0.069)	
Firm Fixed Effects	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Observations	13,077	13,055	13,021	13,045	13,073
R-squared	0.543	0.548	0.543	0.541	0.545

ing status stay negatively significant and the magnitude does not vary either. Meanwhile, the KZ-index seems to have the expected sign: the firms with higher score of KZ-index tend to undertake more wage concessions. The Tobin'Q measure appears to have the same positive sign as the ROA.

In Column (3)-(5), I try different specifications with various combinations for the control variables. One can see the sign and magnitude of the interaction term are stable across all specifications.

3.8.2 Refutability Tests

In this section, I conduct a refutabiliy test to examine if the effect of covenant violation on the change of wage is simply generated by different labor market trends. The idea is that in the absence of big creditors' intervene, firms might not necessarily need to negotiate with workers and thus one would not expect any concession of wage through the pension underfunding channel. In other words, if the DD setup is valid, the interaction term of the time and pension underfunding dummy should not be statistically different from zero. I construct two new dummies variables to conduct the refutability test. The firms will have a covenant violation in the following three years are noted as $Bind_{(0,3)}$ and the firms have had a covenant violation in the past three years are noted as $Bind_{(-3,0)}$. I re-run the baseline regression and present the results in Table 9. One can see from the table that the interaction terms in both specifications are insignificant. It indicates that having a covenants violation within three years' window (either in the next three years or the past three years) is not associated with any change in fimr's wage through the pension underfunding channel.

Table 9: Refutability Tests

This table presents the DID specification with percent change in average wages as the dependent variables. The firm accounting data is from Compustat. The dependent variable from column (1) to (2) is the percent change in average wages. *Bind* is a dummy variable which equals one if firms will have a covenant violation in the next three years (column (1)) or have had a covenant violation in the past three years (column (2)). ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	(1)	(2)
	$\Delta\frac{wages}{employees}$	$\Delta\frac{wages}{employees}$
	$Bind(0,3)$	$Bind(-3,0)$
Bind	-0.024	-0.044
	(0.017)	(0.035)
Underfunding	0.004	-0.001
	(0.004)	(0.003)
Bind*Underfunding	0.081	-0.092
	(0.092)	(0.071)
Asset	0.002	0.003
	(0.004)	(0.003)
Cash Flow	-0.061**	-0.055**
	(0.024)	(0.025)
ROA	-0.151***	-0.146***
	(0.031)	(0.031)
Firm Fixed Effects	Yes	Yes
Year Fixed Effects	Yes	Yes
Observations	12,399	13,021
R-squared	0.541	0.542

Table 10: Loan Covenant Violations, Pension Underfunding and Wages (controlling for industry-year fixed effect)

This table presents the DID specification with percent change in average wages as the dependent variables. The firm accounting data is from Compustat. The dependent variable from column (1) to (5) is the percent change in average wages. Underfunding dummy is an indicator variable that equals one if underfunding is larger than 0% (column 1), 10% (column 2), 15% (column 3), 20% (column 4) and 25% (column 5). Bind is a dummy variable which equals to one if there is a covenant violation in year $t-1$. The regression include an interaction term fixed effect between two-digit SIC code and year. ***,** and * indicate significance at 1%, 5% and 10% levels, respectively. All standard errors are clustered at the firm level.

	Underfunding >				
	0%	**10%**	**15%**	**20%**	**25%**
	(1)	(2)	(3)	(4)	(5)
	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %	$\Delta\frac{wages}{employees}$ %
Bind	-0.024	-0.065***	-0.072***	-0.075***	-0.081***
	(0.017)	(0.016)	(0.016)	(0.016)	(0.015)
Underfunding	0.004	-0.002	-0.005	-0.005	-0.007*
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Bind*Underfunding	**-0.103*** **	**-0.065*** **	**-0.057*** **	**-0.056*** **	**-0.045** **
	(0.021)	**(0.020)**	**(0.022)**	**(0.022)**	**(0.022)**
Asset	0.003	0.004	0.004	0.004	0.004
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Cash Flow	-0.059**	-0.060**	-0.060**	-0.060**	-0.058**
	(0.024)	(0.025)	(0.025)	(0.025)	(0.025)
ROA	-0.147***	-0.151***	-0.151***	-0.152***	-0.154***
	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)
Firm Fixed Effects	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Industry*Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Observations	13,077	13,077	13,077	13,077	13,077
R-squared	0.543	0.542	0.542	0.542	0.542

3.8.3 Industry-Year Fixed Effects

My empirical strategy aims to study within firms wage renegotiation by using firm-level on wage, pension underfunding. While my results still hold after controlling for firm and year fixed effects, it is still possible that the significance of the coefficient captures the degree of industry financial distress over time as opposed to pure within-firm cross-sectional variation in pension fundings. I attempt to relief the concern by including industry*year fixed effects in the regressions.

Results are reported in Table 10. The inclusion of industry*year fixed effects enable me to identify a cleaner within-firm comparison. The coefficients of interaction term in Table 10 are negative and similar to the baseline regression in Table 4 and Table 5. The inclusion of extra industry*year fixed effects diminishes the statistical significance of the results. The difference-in-difference estimator is 6.5% when the pension plan is underfunded by at least 10%, while it is 5.7%, 5.6% and 4.5% when the pension plan is underfunded by at least 15%, 20% and 25%, respectively.

3.9 Conclusion and Policy Implication

In this paper, I show the role of pension plans in shaping the firms' labor market decision. More specifically, by employing the loan covenants violation and consequently transferring of control rights to creditors, I examine the strategic use of pension underfunding by firms and the resultant wage cuts. The study sheds light on how firms strategically renegotiate labor contracts to extract concessions from labor. Also, the evidence suggests that credit contracts between debt-holders and shareholders have spillover effects on non-financial stakeholders, and the funding status of pension plans is among the first-order important channels. I also find that the wage concession is less server for firms from industry with bigger bargaining power. Surprisingly, firms do not

appear to strategically use pension underfunding to renegotiate wages during bankruptcy process.

The literature on DB corporate pension plans has contributed to our understanding of the interactions between the different stakeholders in a firm. It has shed light on the nature of the employment contract between firms and workers and the conflicts of interest that may arise between them, and between sponsoring firms and the government, managers and shareholders, among others. My study has at least two folds of policy implications. First, given that the employees has very weak protections over their pension claims towards firms, policy makers should think of ways to motivate firms to have a better pension plan risk management, particularly for firms with DB pension benefits. Second and more broadly speaking, in order to reduce firms' incentive to 'dump' pensions during period of financial downturn, policy maker should come up with a better institution to govern the complex pension plan market.

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Table A1: Maximum Annual Guarantees Given by PBGC

This table provides the maximum annual amount covered by the Pension Benefit Guaranty Corporation gives at different ages at the time of retiring. Data source: `http://www.pbgc.gov/`.

	50 Years	55 Years	60 Years	65 Years
1992	$9,879	$12,702	$18,348	$28,227
1993	$10,238	$13,163	$19,013	$29,250
1994	$10,739	$13,807	$19,943	$30,682
1995	$10,810	$13,899	$20,076	$30,886
1996	$11,097	$14,267	$20,608	$31,705
1997	$11,598	$14,911	$21,539	$33,136
1998	$12,099	$15,556	$22,469	$34,568
1999	$12,815	$16,476	$23,769	$36,614
2000	$13,531	$17,397	$25,128	$38,659
2001	$14,247	$18,317	$26,458	$40,705
2002	$15,034	$19,330	$27,921	$42,955
2003	$15,392	$19,790	$28,585	$43,977
2004	$15,535	$19,974	$28,851	$44,386
2005	$15,965	$20,526	$29,649	$45,614
2006	$16,681	$21,447	$30,978	$47,659
2007	$17,316	$22,272	$32,172	$49,500
2008	$18,108	$23,280	$33,636	$51,744
2009	$18,900	$24,300	$35,100	$54,000
2010	$18,900	$24,300	$35,100	$54,000
2011	$18,900	$24,300	$35,100	$54,000
2012	$19,536	$25,128	$36,288	$55,836
2013	$20,112	$25,860	$37,356	$57,468
2014	$20,760	$26,688	$38,556	$59,316
2015	$21,036	$27,060	$39,084	$60,132

Figure 1: Loan Covenant Violations, Pension Underfunding and Wage

The sample consists of 13077 firm-year observations that are bound by either a current ratio or a net worth covenant during the period 1994 to 2010. This figure shows percentage annual changes in the average wage in event time leading to and after the year when a covenant violation occurs. 1st (4th) quantile of underfunding is the first (fourth) quantile of the difference between pension plan liability and pension plan asset in the previous fiscal year (year $t-1$).

Figure 3.1

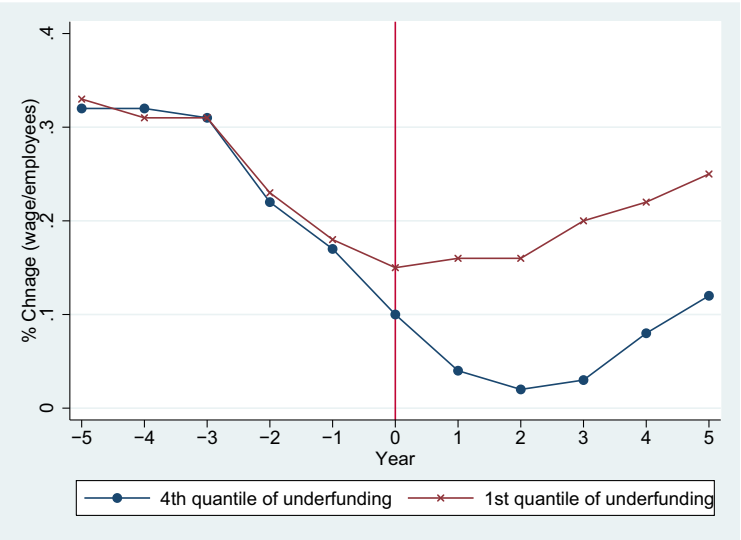

Figure 2: Loan Covenant Violations, Pension Underfunding and Other Firm Characteristics

The sample consists of 13077 firm-year observations that are bound by either a current ratio or a net worth covenant during the period 1994 to 2010. This figures show firm's financial characteristics (cash flows, ROA, investment and leverage) in event time leading to and after the year when a covenant violation occurs. 1st (4th) quantile of underfunding is the first (fourth) quantile of the difference between pension plan liability and pension plan asset in the previous fiscal year (year $t-1$).

Figure 3.2