The China Syndrome: Local Labor Market Effects of Import Competition in the United States: Comment*

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Abstract

We re-examine the findings by Autor, Dorn, and Hanson (ADH, *American Economic Review 2013, 103(6)*) on the impact of Chinese import penetration on U.S. local employment by taking into account the concurrent housing boom. The responses of total employment, unemployment, or not-in-the-labor force to import exposure fall by about one-half when controlling for changes in housing prices, and become statistically insignificant in a number of cases. Results across sectors are more subtle. Noncollege workers in the manufacturing sector continue to experience a reduction in employment after correcting for the 'masking' effect of the housing boom, but that reduction does not occur in the nonmanufacturing sector. For college workers, their employment in the nonmanufacturing sector even rises due to the China shock, which fully offsets their reduced employment in manufacturing during 2000-2007. Our results imply that the net reduction in total US employment due to Chinese import exposure was about 0.8 million workers, or less than one-half of that implied by the estimates in ADH (2013).

JEL Codes: F14, F16, J23, R23

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1 Introduction

In an influential study, Autor, Dorn, and Hanson (2013, hereafter ADH 2013) show that rising import competition from China has been an important contributor to the recent decline in the employment rate of working age population in the United States. Exploiting variation in exposure to Chinese import across local labor markets (commuting zones) over 1990-2007, they find that Chinese import exposure caused a large reduction in manufacturing employment: a \$1,000 per worker increase in import exposure over a decade reduces manufacturing employment per working-age population by 0.6 percentage points (their Table 3, column 6), explaining about 44 percent of the actual decline in manufacturing employment from 1990 through 2007. Furthermore, the negative employment shock by Chinese imports goes beyond manufacturing and exists for nonmanufacturing workers. To be more specific, import exposure to Chinese imports caused a substantial employment decline in *both* manufacturing and nonmanufacturing sectors for workers without college education; while for workers with college education, import exposure caused substantial job loss in manufacturing sectors but a statistically insignificant increase in employment in nonmanufacturing sectors (their Table 5, Panel B).

Such results, accompanied with findings in other studies such as Pierce and Schott (2015) and Acemoglu et al. (2016), have important policy implications and challenge the benign view towards globalization. In particular, ADH (2013) show that for regions (i.e. commuting zones) facing competition from China, the reduction in manufacturing jobs has not been offset by rising employment in nonmanufacturing sectors. For noncollege workers, the displacement effect in nonmanufacturing sectors is at a similar magnitude as that in the manufacturing sector. These results suggest either a lack of reallocation between sectors and/or a large negative aggregate demand effect at the local level due to competition from China.

In this note, we re-examine the data and show that different conclusions can be reached if we also take into account the concurrent impact of the housing boom. The response of the total employment-to-population rate to import exposure falls by about one-half and becomes statistically insignificant in a specification that includes changes in local housing prices. Similar effects are found for the unemployment rate and the fraction of the population not in the labor force. These results indicate that during the same period when manufacturing workers faced import competition, nonmanufacturing sectors – in particular the construction sector – experienced a demand boom. This idea has been elaborated in Charles, Hurst, and Notowidigdo (2016a, 2016b), who argue that during 2000-2007 unemployment was 'masked' by the strong housing boom which promoted employment in the construction sector. In the case of ADH (2013), if commuting zones that experienced larger changes in import exposure also had smaller increases in housing prices, then omitting any variables for housing demand would bias upwards (in absolute value) the estimated effect of import exposure. We show that this is indeed the case: there is a negative correlation between Chinese import exposure and housing prices. Furthermore, we show that such effect goes beyond a 'masking' effect that mainly affects workers without college education (Charles et al. 2016b). Even for college educated workers, after controlling for housing prices changes, the negative impact of import exposure on overall employment drops by more than one-half as a result of rising employment in nonmanufacturing sectors.

Given the fact that there was a housing boom during 2000-2007, while at the same time imports from China started to accelerate after 2001 when China gained accession into the WTO, we also explore splitting the sample into two periods (that is: 1990-2000 and 2000-2007) instead of pooling them. We find that for the 1990-2000 period, import exposure to Chinese imports did not impose a significant impact on US employment (therefore justifing the focus on wage structure by the previous literature) in either the manufacturing or nonmanufacturing sectors, for both college and college educated workers. Chinese imports did cause a substantial reduction in the manufacturing employment rate in the 2000-2007 period. But for college workers, this reduction was *fully* offset by rising employment in the nonmanufacturing sector. Within the ADH (2013) data, therefore, is a source of employment gains from the China shock that becomes more apparent when correcting for the housing boom, and which should be explored further.

The rest of the paper is structured as follows: Section 2 presents the results when controlling for housing prices. Section 3 confirms that these results continue to hold when treating housing prices as endogenous. Several robustness checks are in section 4, and section 5 concludes.

2 The Role of the Housing Boom

ADH (2013) point out that US regions (commuting zones) have different exposure to import competition from China due to their industry structure. Those regions that have larger shares of employment in industries that experienced larger growth of Chinese imports (at the national level) will suffer more from import competition from China. Their specification is:

$$\Delta L_{it} = \gamma_t + \beta_1 \Delta IPW_{uit} + X_{it}\beta_2 + \delta_r + e_{it}, \tag{1}$$

where ΔL_{it} is the decadal change (that is, 2000 minus 1990, or 2007 minus 2000) in the employment share of the working-age population in commuting zone i. In different specifications, we distinguish ΔL_{it} by sectors (manufacturing vs. nonmanufacturing), and/or by education level (college vs. noncollege), or replace it with unemployment rate or the fraction of the population not in the labor force (NILF). ΔIPW_{uit} is the change in import exposure, which is instrumented by China's exports to other high-income countries.[1] γ_t is a time dummy for each decadal period. The vector X_{it} contains a set of economic and demographic controls at the start of each decade.[2] In addition, δ_r augments the model with geographic dummies for the nine Census divisions to absorb region-specific trends.

During the same period when imports from China grew quickly, the US also experienced large changes in housing price. Importantly, such changes in housing price also vary across regions. For example, Charles, Hurst, and Notowidigdo (2016b) argue that during the same period when there was large and persistent decline in manufacturing employment, in 2000-2006, the housing boom simultaneously increased employment in construction. Thus, the housing boom 'masks' the adverse labor market effects of the manufacturing decline. In addition, the housing boom may also affect employment through the collateral channel, whereby firms that own real estate increase

[1]The measurement of ΔIPW_{uit} by ADH (2013, p. 2128) uses the change in US imports from China in each industry, weighted by initial commuting zone employment relative to total US employment in each industry, and summed over industries leading to the additional subscript u for the US. The idea for the instrument is that China's exports to other high-income countries are correlated with the US imports from China due to productivity improvements in China or falling trade costs within the same sector. ADH (2013) use the China's total exports to eight countries: Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland. The change in these exports by industry are weighted by decade-old commuting zone employment relative to total US employment in each industry, and summed over industries, to obtain the instrument.

[2]These control variables include: the percentage of employment in manufacturing, percentage of college-educated population, percentage of foreign-born population, percentage of employment among women, percentage of employment in routine occupations, and the average offshorability index of occupations.

their investment in response to rising real estate prices (Chaney, Sraer, and Thesmar, 2012). So, if commuting zones that experienced larger changes in import exposure also had smaller increases in housing prices, then omitting any variable for housing demand would bias up (in absolute value) the estimated effect of import exposure.

To explore whether this bias occurs, we augment ADH's original specification in equation (1) by including a measure of housing price changes at the commuting zone level, resulting in a balanced panel of 522 commuting zones with housing information.[3] Table 1 compares the descriptive statistics of key variables of interest between the complete sample (used by ADH, 2013) and the matched sample with housing price data (used in this paper). The sample with housing price data accounts for 97.7 percent of the US population, and it closely resembles the original, complete sample in the statistics of key variables.

Table 1: Summary Statistics

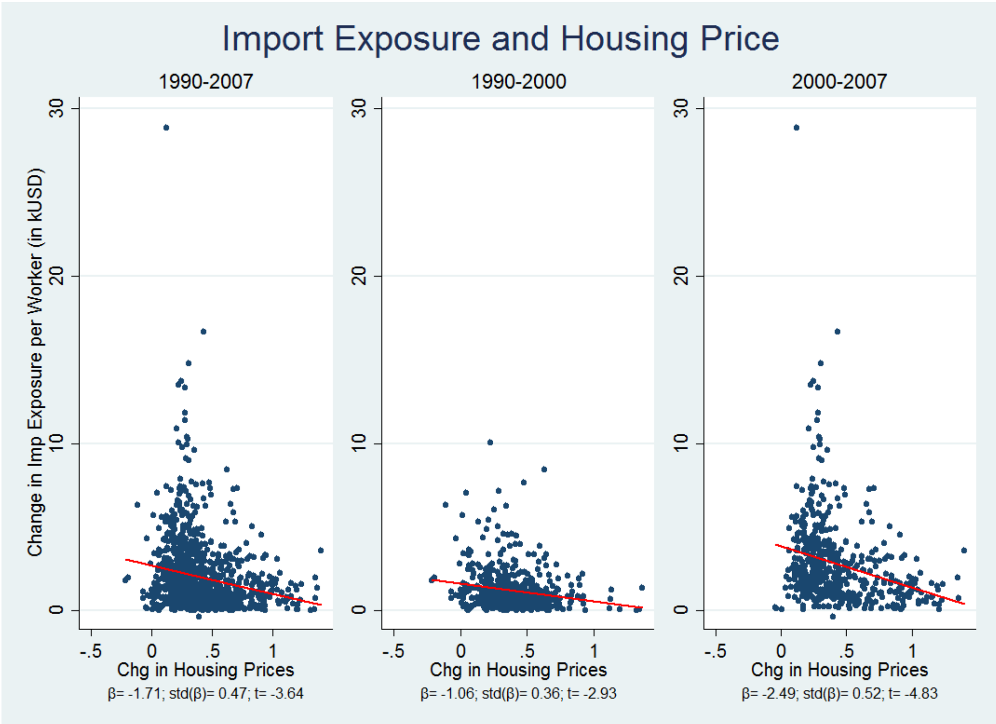
Variable	Mean	Std.Dev
Full sample (N=1,444 pooling 1990-2000 and 2000-2007)		
Δ Imports from China/workers	1.884	1.752
Δ manuf. employment/working-age pop	-2.401	1.746
Δ non-manuf. employment/working-age pop	2.496	2.819
Matched sample with housing price data (N=1,044, population share = 97.7%)		
Δ Imports from China/workers	1.869	1.668
Δ manuf. employment/working-age pop	-2.417	1.707
Δ non-manuf. employment/working-age pop	2.483	2.817

Note: The full sample (N = 1,444 = 722 commuting zones × 2 time periods) is adopted from the data in ADH (2013). The matched sample (N = 1,044 = 522 commuting zones × 2 time periods) includes commuting zones that have housing price information.

Figure 1 shows the correlation between changes in local import exposure from China and changes in the local housing price index, for the pooled period 1990-2007, and for 1990-2000 and 2000-2007 separately. There is an obvious negative correlation between the two variables of interest, which appears to be much stronger in the 2000-2007 period. For the later period, regions that experienced larger import shocks also experienced smaller increases in housing prices.

[3] The annual Housing Price Index at U.S. county level can be downloaded from the Federal Housing Finance Agency website: https://www.fhfa.gov/DataTools/Downloads/pages/house-price-index.aspx. After aggregating and matching country to commuting zones, we have housing price information for 522 commuting zones in 1990-2007. The excluded commuting zones are mainly rural regions without housing price information.

Figure 1: Correlation of import exposure and housing price changes



Note: This figure shows the correlation between changes in import exposure and changes in housing price index, across 522 commuting zones.

Table 2 reports the results of the augmented regressions. In all cases, imports from China are instrumented by Chinese exports to other advanced economies, the method proposed by ADH to capture the supply shock of Chinese exports. For ease of comparison, Panel I reproduces ADH's results using the two decadal periods stacked, with 722 commuting zones. Panel II then uses the matched sample of commuting zones with housing price data (i.e, 522 commuting zones). The dependent variables, as indicated in the top of each column, are the manufacturing employment-to-population rate (column 1), nonmanufacturing employment-to-population rate (column 2), total employment-to-population rate (column 3), unemployment rate (column 4), and not-in-the-labor-force rate (NILF, column 5). By the definition of the population shares and the property of linear regressions, we know that the coefficents for each column satisfy the relationship $col1 + col2 = col3 = -(col4 + col5)$.

Panel II confirms the results of ADH: import exposure substantially reduces the manufacturing employment rate while it has a negative but insignificant effect on the nonmanufacturing employment rate, resulting in rising unemployment rate and NILF rate. In fact, the effect of China import

Table 2: Imports from China and US Employment

Dependent variables: Changes in population shares by employment status

	(1) Mfg emp	(2) Non-mfg emp	(3) Total Emp	(4) Unemp	(5) NILF
Panel I: ADH sample, 722 CZ					
All education levels					
(Δ imports from China) /worker	-0.596***	-0.178	-0.774***	0.221***	0.553***
	(0.099)	(0.137)	(0.176)	(0.058)	(0.150)
College education					
(Δ imports from China) /worker	-0.592***	0.168	-0.424***	0.119***	0.304***
	(0.125)	(0.122)	(0.123)	(0.039)	(0.113)
No college education					
(Δ imports from China) /worker	-0.581***	-0.531***	-1.112***	0.282***	0.831***
	(0.095)	(0.203)	(0.252)	(0.085)	(0.211)
Panel II: Matched sample, 522 CZ					
All education levels					
(Δ imports from China) /worker	-0.661***	-0.177	-0.837***	0.240***	0.597***
	(0.100)	(0.161)	(0.198)	(0.062)	(0.175)
College education					
(Δ imports from China) /worker	-0.651***	0.221*	-0.431***	0.129***	0.302**
	(0.134)	(0.134)	(0.139)	(0.044)	(0.129)
No college education					
(Δ imports from China) /worker	-0.646***	-0.581**	-1.228***	0.305***	0.922***
	(0.098)	(0.238)	(0.284)	(0.094)	(0.247)
Panel III: Matched sample, controlling for changes in housing prices					
All education levels					
(Δ imports from China) /worker	-0.529***	0.180	-0.349	0.122**	0.227
	(0.079)	(0.202)	(0.212)	(0.057)	(0.204)
Δ housing price index	1.722***	4.654***	6.376***	-1.549***	-4.827***
	(0.384)	(0.676)	(0.980)	(0.416)	(0.779)
College education					
(Δ imports from China) /worker	-0.514***	0.407***	-0.107	0.046	0.061
	(0.112)	(0.145)	(0.128)	(0.047)	(0.122)
Δ housing price index	1.799***	2.426***	4.225***	-1.079***	-3.146***
	(0.408)	(0.241)	(0.398)	(0.320)	(0.335)
No college education					
(Δ imports from China) /worker	-0.519***	-0.056	-0.575*	0.148*	0.427
	(0.092)	(0.279)	(0.323)	(0.086)	(0.301)
Δ housing price index	1.662***	6.851***	8.513***	-2.050***	-6.463***
	(0.409)	(1.173)	(1.500)	(0.561)	(1.161)

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
The ADH sample has 722 commuting zones in each period, while the matched sample has 522 commuting zones in each period for which we have housing price data. All regressions contain a dummy for the 2000-2007 period, a set of census division dummies, and the full set of control variables for the start of period economic and demographic conditions, which includes percentage of employment in manufacturing, percentage of college-educated population, percentage of foreign-born population, percentage of employment among women, percentage of employment in routine occupations, and finally average offshorability index of occupations. All regressions are weighted by start of period commuting zone's share of national population.

exposure is even a bit stronger for the reduced sample of commuting zones: a $1,000 per worker increase in a CZ's import exposure reduces its manufacturing employment to population rate by 0.66 percentage points, and its nonmanufacturing employment rate by 0.18 percentage points (though not significant), resulting in a total drop in the employment rate of 0.84 percentage points. Further decomposing the effect by workers' education levels confirms what ADH find: a $1,000 import exposure from China reduces the noncollege employment rate by a highly significant 1.23 percentage points, consisting of 0.65 percentage points drop in manufacturing and 0.58 percentage points drop in nonmanufacturing sector, respectively. College employment in manufacturing also experienced a significant drop of 0.65 percentage points, while in nonmanufacturing it increases by 0.22 percentage points, with the net negative effect on employment rate at a highly significantly 0.43 percentage points.

The effect of import exposure, however, is mitigated if we augment the ADH specification with local housing price changes, a variable we use to capture the local housing demand shocks. The results are reported in Panel III. First, the effects on both the employment rate (column 3) and the unemployment rate (column 4) drop by about one-half, regardless whether we consider the combined sample with all education levels or we separate college employment and noncollege employment. For example, a $1,000 increase in Chinese imports per worker reduces employment by 0.84 percentage points in Panel II, but when we include housing price changes, the reduction drops to 0.35 percentage points in Panel III. Second, in a number of cases, the effect of import exposure on total employment, unemployment or NILF loses its statistical significance in Panel III. Thirdly, in all cases, local housing booms promote employment in both manufacturing and nonmanufacturing strongly and significantly, and reduce unemployment and the share of NILF. The impact of import exposure on overall *manufacturing* employment drops by 20 percent, from 0.66 to 0.53 percentage points, but remains significant when we include the housing price index in the regressions.

Strikingly, some conclusions by ADH no longer arise. In particular, ADH find that import exposure reduced the noncollege employment rate, at similar magnitudes in both the manufacturing and nonmanufacturing sectors. Including the changes in housing demand, however, makes the impact of import exposure on noncollege employment in the nonmanufacturing sector much smaller in magnitude and no longer significant. The net employment effect on noncollege workers is a 0.58 percentage point drop for each $1,000 increase in Chinese imports per worker, which is about one-half of the effect when we do not consider housing, indicating the importance of correcting for the 'masking' effect of the housing boom. As for college workers, ADH (2013) find that import exposure reduced their employment rate substantially in manufacturing, but had only a modest and insignificant effect in the nonmanufacturing sector. In contrast, we find that the detrimental effect on college workers in manufacturing is still there, but that the effect on those in nonmanufacturing turns *positive and significant* after including housing price changes. The effects on manufacturing and nonmanufacturing offset each other, resulting in an insignificant impact of import exposure on total employment for college workers.

We can apply the observed import growth from China to quantify the economic magnitude of import exposure on manufacturing and nonmanufacturing workers. According to ADH (2013), Chinese import exposure rose by $1,140 per worker between 1990 and 2000 and by an additional $1,839 per worker between 2000 and 2007. Using their estimated coefficients (Table 2, Panel I), they find a net reduction in US manufacturing employment of 1.53 million workers for the full period 1990-2007. Applying the same calculation to nonmanufacturing employment implies another 0.46 million job loss, or 1.99 million lost jobs in total.

A similar calculation, but using the coefficients that we re-estimated controlling for housing price changes (Table 2, Panel III), shows that although Chinese import exposure caused a reduction in manufacturing employment by 1.33 million workers, it also caused an *increase* in nonmanufacturing employment by 0.50 million. Thus the net reduction in total US employment due to rising Chinese import exposure was about 0.83 million workers, or less than one-half of that implied by the estimates in ADH (2013). This striking contrast comes mainly from the different response from the nonmanufacturing sector. In particular, rising Chinese import exposure led to a negligible reduction in noncollege nonmanufacturing jobs (0.066 million), while it increased employment for college nonmanufacturing workers by 0.56 million.[4]

[4] See the Appendix and note 12 for the details of these calculations.

3 Instrumental Variable Estimation

One concern about controlling for the effect of housing boom is that the changes in local housing price may be the result of import exposure, at least in part. Feler and Senses (2016), for example, show that import shocks depress the self-reported home values in U.S. commuting zones. In this case, our estimation in Table 2 may understate the effect of import exposure. Other factors can also lead to endogeneity, too. On the one hand, unobserved local conditions may affect employment and housing prices simultaneously. On the other hand, local job opportunities can also reversely affect housing prices.

To deal with these endogeneity issues, Table 3 instruments for the housing price changes using the rapid changes in housing prices that occurred in the local area (Ferreira and Gyourko, 2011; Charles et al. 2016b). The idea is that underlying fundamentals for housing demand (such as productivity, income, or population) do not change abruptly and are smoothly incorporated into prices when they do change, so that sharp breaks from the trend arguably reflect variations due to exogenous speculative activities or other housing-specific forces.

More specifically, following Charles et al. (2016b), we estimate for each local area an OLS regression with a structural break, and search for the break date that maximizes the R^2 of the regression:

$$lnP_{it} = \omega_i + \tau_i t + \lambda_i (t - t_i^*) D_{it} + \epsilon_{it}, \quad (2)$$

where lnP_{it} is the log value of quarterly housing price index for each area i in year-quarter t, and ω_i is a constant. D_{it} is a dummy variable which equals 1 for periods after the date of structural break t_i^*, and 0 otherwise. Thus τ_i is the linear time trend before the structural break, while λ_i captures the size of the structural break. Our estimation is run for each metropolitan statistical area (MSA) for which the quarterly housing price series is available. We then match MSAs with commuting zones, which reduces the sample to 321 commuting zones during 2000-2007 and 291 commuting zones during 1990-2000.

We estimate equation (2) for each local area separately over periods 1990-2000 and 2000-2007, and use the annualized size of the structural break λ_i as the instrument for the decadal changes in housing prices.[5] Since most areas experienced housing booms in the second period while they had a slowdown in the first period, the first stage coefficients for the structural break λ_i in the two periods have opposite sign. To capture this difference, we interact λ_i with the period dummies to instrument *separately* for the housing price changes in the two periods.[6]

Table 3: Imports from China and US Employment: Instrumenting Housing Prices

	(1) Mfg emp	(2) Non-Mfg emp	(3) Total emp	(4) Unemp	(5) NILF
	Panel I: All education levels				
(Δ imports from China)/worker	-0.577***	0.219	-0.358	0.201***	0.157
	(0.086)	(0.236)	(0.234)	(0.070)	(0.232)
Δ housing price index	1.518***	5.189***	6.707***	-1.324**	-5.384***
	(0.480)	(1.189)	(1.542)	(0.537)	(1.246)
	Panel II: College education				
(Δ imports from China)/worker	-0.567***	0.479***	-0.088	0.120***	-0.032
	(0.133)	(0.175)	(0.152)	(0.047)	(0.140)
Δ housing price index	1.452***	3.477***	4.929***	-0.890**	-4.039***
	(0.500)	(0.340)	(0.584)	(0.385)	(0.432)
	Panel III: No college education				
(Δ imports from China)/worker	-0.548***	-0.006	-0.554	0.232**	0.322
	(0.098)	(0.343)	(0.379)	(0.109)	(0.371)
Δ housing price index	1.830***	7.262***	9.091***	-1.845**	-7.246***
	(0.566)	(2.133)	(2.560)	(0.745)	(2.090)

	First Stage Results	
	(1) (Δ imports from China)/worker	(2) Δ housing price index
(Δ imports from China to Other)/worker	0.569***	-0.022**
	(0.090)	(0.009)
Structural break in housing price 1990-2000	-1.800	-2.433***
	(1.164)	(0.334)
Structural break in housing price 2000-2007	-2.229	3.385***
	(1.531)	(0.467)
First Stage F statistics	13.55	91.18

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
The MSA sample includes 321 commuting zones during 2000-2007 and 291 commuting zones during 1990-2000 that we have matched with housing prices data and the instruments for housing price changes, both of which are at the level of the metropolitan statistical area. All regressions include the full set of control variables from Table 2. All regressions are weighted by start of period commuting zone's share of national population. The bottom panel reports the first stage results for the instruments and the F statistics. The joint Kleibergen-Paap Wald rk F statistic is 13.67.

[5]For the first period instrument, we use the quarterly price series between 1990Q1 and 2000Q4 while restricting the break date to be between 1991Q1 to 2000Q4. For the second period instrument, we use the quarterly price series between 2000Q1 and 2005Q4 while restricting the break date to be between 2001Q1 and 2005Q4. We restrict the break to occur before 2006 since housing booms had already started to burst in 2006 for some MSAs. Extending the search of the break date to 2006Q4, however, leads to qualitatively similar results.

[6]We plot the estimated dates and sizes of the structural breaks in Appendix Figure A.1. We show that during the first period there is a negative correlation between housing price changes and the size of the breaks, while housing price changes in the second period are clearly positively correlated with the size of the structural breaks. In Appendix Figure A.2, we provide a few city examples for the housing price index and the structural break, illustrating the slowdown in housing prices in the first period and the boom in the second period.

Panel I of Table 3 reports the results with all education levels: a \$1,000 increase in import exposure per worker leads to a 0.58 percentage point reduction in manufacturing employment, while at the same time a 0.22 percentage point increase in nonmanufacturing employment. Panels II and III look at college and noncollege employment respectively. Consistent with the results reported in Table 2, although import exposure substantially reduced manufacturing employment for both college and noncollege workers, it also created a strong increase in employment for college workers in the nonmanufacturing sector, and it has no significant effect on noncollege worker employment in the nonmanufacturing sector.

In the bottom Panel, we report the first stage regressions. As expected, they show a strong negative effect of the first period structural break and a strong positive effect of the second period structural break on decadal housing price changes. Interestingly, the instrument for import exposure, which is Chinese import exposure to other countries, has a slight and negative impact on housing price changes, implying import exposure does reduce house values. Encouragingly, the coefficient estimates for import exposure in Tables 2 and 3 are very similar, with or without instrumenting for housing price changes, despite the fact that Table 3 has the reduced MSA sample.[7] If we re-do the calculation for the total number of jobs lost due to the Chinese import exposure with coefficients in Table 3, we find that the net reduction in total US employment due to rising Chinese import exposure was about 0.78 million workers (including a reduction of 1.43 million manufacturing jobs and an increase of 0.65 million nonmanufacturing jobs), which is very similar to the results implied by Table 2.

4 Robustness Checks

In the previous section, we stacked the differences for the two periods (i.e. 1990-2000 and 2000-2007), as in ADH. However, China's exports to the US started to accelerate after 2001, when China gained its accession into the WTO. Before 2001, the share of US imports from China increased by about 0.5 percentage point annually, while after 2001 it jumped to growing by 1.2 percentage points per year. In terms of import penetration by China (i.e., imports from China as a share of

[7]To compare, we report in the Appendix the ADH results using the same reduced sample (Table A.1), without controlling for housing prices changes, and the second stage regressions as in Table 3 but without instrumenting for housing price changes (Tables A.2).

total US expenditure), that increased by 0.14 percentage points per year before 2001 and then by 0.38 percentage points after 2001.

Moreover, the literature has attributed the dynamics in labor market outcome for the period 1990-2000 mainly to skill biased technology change which substitute for routine jobs (Autor, Katz, and Kearney, 2008), or to service offshoring (Crinò, 2010), instead of import competition. If there is any structural difference in labor market response to import exposure, focusing on the second period (2000-2007) should give even stronger results. Table 4 reports the results when we split the sample into two periods. Panel I considers workers at all education levels, while Panels II and III consider workers with or without college education, respectively. In all cases, we also control for the effect of housing price changes.

The results in Panel I show that in the 1990s, imports from China did not exert a significant impact on total employment, in either the manufacturing or the nonmanufacturing sectors. This result is probably not surprising, given that China imports had not yet taken a substantial share in the US market. For the later period, 2000-2007, import exposure from China substantially reduced the manufacturing employment share, which is partly offset by the positive effect on nonmanufacturing employment. The net effect on the employment rate is negative, but not significant.

To check the statistical equality of coefficients in the regressions for two different periods, we run a set of Chow tests. The joint equivalence of coefficicents for import exposure, housing price, and all control variables X_{it} in (1)[8] are always rejected with near-zero p-values. The Chow tests on the joint equivalence of coefficicents for the two key variables (i.e., import exposure and housing prices) are reported at the bottom of each panel in Table 4. In most cases, we cannot reject the null hypothesis of equivalence of the two coefficients between two regressions. For the college employment share in both manufacturing and nonmanufacturing sectors, however, the p-values for the Chow test is about 0.05, therefore rejecting the equal coefficients across the two separate regressions.

Panel II looks at import exposure and employment status by workers with college education. The results show striking differences in the employment response to import exposure during the 1990s versus the 2000s. In particular, the effect of import exposure on college workers turn out to

[8]See footnote 2.

Table 4: Imports from China and US Employment: Splitting the Sample

Dependent variables: Changes in population shares by employment status

	(1) Mfg emp	(2) Non-mfg emp	(3) Total Emp	(4) Unemp	(5) NILF
Panel I: All education levels					
Sample Period: 1990-2000					
(Δ imports from China) /worker	-0.216	0.332	0.116	-0.117	0.001
	(0.226)	(0.253)	(0.252)	(0.090)	(0.207)
Δ housing price index	0.977*	1.938***	2.915***	-1.381***	-1.535***
	(0.551)	(0.577)	(0.747)	(0.284)	(0.596)
Sample Period: 2000-2007					
(Δ imports from China) /worker	-0.514***	0.308**	-0.206	0.111	0.095
	(0.128)	(0.139)	(0.162)	(0.110)	(0.099)
Δ housing price index	0.715*	1.391**	2.105***	-0.748	-1.357***
	(0.402)	(0.596)	(0.791)	(0.566)	(0.357)
Chow Test (p-value)	0.474	0.826	0.525	0.219	0.888
Panel II: College education					
Sample Period: 1990-2000					
(Δ imports from China) /worker	0.021	0.304	0.325	-0.074	-0.251
	(0.285)	(0.230)	(0.208)	(0.066)	(0.201)
Δ housing price index	1.491**	-0.098	1.394***	-1.466***	0.072
	(0.605)	(0.630)	(0.435)	(0.210)	(0.431)
Sample Period: 2000-2007					
(Δ imports from China) /worker	-0.542***	0.662***	0.119	0.058	-0.177*
	(0.157)	(0.187)	(0.120)	(0.070)	(0.100)
Δ housing price index	0.461	1.327***	1.789***	-0.790*	-0.999***
	(0.376)	(0.446)	(0.548)	(0.403)	(0.329)
Chow Test (p-value)	0.044	0.058	0.391	0.189	0.133
Panel III: No college education					
Sample Period: 1990-2000					
(Δ imports from China) /worker	-0.278	0.295	0.017	-0.136	0.119
	(0.229)	(0.276)	(0.325)	(0.127)	(0.258)
Δ housing price index	0.328	3.002***	3.330***	-1.068***	-2.262***
	(0.559)	(0.678)	(0.949)	(0.413)	(0.747)
Sample Period: 2000-2007					
(Δ imports from China) /worker	-0.473***	0.028	-0.445*	0.110	0.335**
	(0.122)	(0.214)	(0.254)	(0.157)	(0.164)
Δ housing price index	0.957*	1.953**	2.911***	-0.766	-2.145***
	(0.560)	(0.853)	(1.015)	(0.705)	(0.561)
Chow Test (p-value)	0.554	0.536	0.601	0.537	0.735

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Each period has 522 commuting zones. All regressions contain the same set of controls as Table 2. All regressions are weighted by start of period commuting zone's share of national population. At the bottom of each Panel, we report the p-values for Chow tests on the joint equality of the coefficients on import exposure and the housing price index.

be bigger in the second period than in the first period. This effect is negative and larger in absolute value for the manufacturing sectors, while it is positive and larger for the nonmanufacturing sectors. Contrary to what ADH find, during 2000-2007 the reduction in college employment in manufacturing is *fully offset* by the increase in nonmanufacturing employment. As a result, import exposure from China did not cause unemployment for college workers.

Panel III then shows that import exposure also has a different impact on workers without college education in the two periods. In the 1990s, the effect is relatively small and insignificant. During this period, import exposure did reduce manufacturing employment, but the negative impact is offset by increase in employment in the nonmanufacturing sectors, and none are significant. In the later period, 2000-2007, import exposure has a substantial and negative impact on noncollege employment in manufacturing. And in contrast to ADH (2013), import exposure does not have significant impact on noncollege employment rate in nonmanufacturing sectors, re-confirming the importance of correcting for the 'masking' effect as we discussed before.

One alternative way to examine the 'masking' effect of local housing boom on import shock is to check whether commuting zones with large increases in housing demand responded less to import exposure. So we define regions with large increases in housing prices (top tercile) as booming areas. Following Charles et al. (2016b), we include an interaction term between import exposure and booming areas in the regression:

$$\Delta L_{it} = \gamma_t + \beta_1 \Delta IPW_{uit} + \beta_2 \Delta IPW_{uit} \times \{boomingCZ = 1\} + X_{it-1}\beta_3 + e_k. \qquad (3)$$

Now if housing booms 'mask' the real effect of import competition, we would expect β_2 to be significant and opposite in sign to β_1. As before, we instrument for ΔIPW_{uit} using the contemporaneous growth of Chinese exports to other high-income countries, and we also interact that instrument with the indicator variable for booming areas.

The results are shown in Table 5. Panel I stacks the two periods using the matched sample with all education levels. It shows that import exposure from China did cause a loss of jobs in both the manufacturing and nonmanufacturing sectors. In commuting zones with housing booms, however, the negative effect of imports is attenuated in manufacturing and even offset in the nonmanufacturing sector. The housing effects become clearer when we separately look at employment by education levels. Panel II reports college employment. In the top one-third of commuting zones with increasing housing prices, import exposure causes a reduction in manufacturing employment but a rise in nonmanufacturing employment, so there is virtually no effect at all on total employment. Panel III presents the case for noncollege workers. In housing boom areas, import competition still reduces manufacturing employment, but to a lesser extent than in areas without a housing boom. Furthermore, the reduction in nonmanufacturing employment is completely offset by the interaction term, so there is no net impact of the China shock on nonmanufacturing employment of workers without a college education.

Table 5: Imports from China and US Employment, with Housing Boom Interaction

Dependent variables: Changes in population shares by employment status

	(1) Manuf. emp	(2) Non-mfg emp	(3) Total emp	(4) Unemp	(5) NILF
Panel I: All education level					
(Δ imports from China)/worker	-0.714***	-0.368**	-1.082***	0.288***	0.794***
	(0.128)	(0.170)	(0.273)	(0.084)	(0.213)
Δ import exposure $\times$ top 1/3 housing boom	0.194*	0.690***	0.884***	-0.173*	-0.711***
	(0.104)	(0.234)	(0.295)	(0.099)	(0.233)
Panel II: College education					
(Δ imports from China)/worker	-0.695***	0.097	-0.598***	0.160***	0.439***
	(0.155)	(0.097)	(0.184)	(0.052)	(0.161)
Δ import exposure $\times$ top 1/3 housing boom	0.156	0.447***	0.604***	-0.111	-0.493***
	(0.106)	(0.134)	(0.159)	(0.073)	(0.113)
Panel III: No college education					
(Δ imports from China)/worker	-0.715***	-0.853***	-1.568***	0.373***	1.194***
	(0.120)	(0.294)	(0.373)	(0.121)	(0.284)
Δ import exposure $\times$ top 1/3 housing boom	0.246**	0.980***	1.226***	-0.245*	-0.981***
	(0.122)	(0.349)	(0.445)	(0.135)	(0.359)

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Each period has 522 commuting zones. All regressions include the full set of control variables from Table 2. All regressions are weighted by start of period commuting zone's share of national population.

In Table 6 ,we consider the impact of import exposure on average weekly wage of US workers. Similar to Table 2, Panel I replicates the results of ADH (2013, Table 6), while Panel II uses the matched sample of 522 commuting zones. Both panels show a significant negative effect of import exposure on the average wages of workers within CZs (column 1). Such negative effects mainly show up for nonmanufacturing workers, while the effect on manufacturing workers' wages is not significant. The same pattern holds when we separately look at college and noncollege workers. But when Panel III includes changes in housing prices, all depressing effects of import exposure

Table 6: Imports from China and US Wages

Dependent variables: Ten-year equivalent change in average log weekly wage (in log pts)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
		All Workers			College education			No college education	
	All	Manuf.	Non-manuf.	All	Manuf.	Non-manuf.	All	Manuf.	Non-manuf.
				Panel I: ADH sample, 722 CZ					
(Δ imports from China)/worker	-0.759***	0.151	-0.761***	-0.757**	0.458	-0.743**	-0.814***	-0.101	-0.822***
	(0.253)	(0.482)	(0.261)	(0.308)	(0.340)	(0.297)	(0.236)	(0.369)	(0.246)
				Panel II: Matched sample, 522 CZ					
(Δ imports from China)/worker	-0.806***	0.156	-0.791**	-0.824**	0.508	-0.812**	-0.844***	-0.140	-0.813***
	(0.295)	(0.561)	(0.307)	(0.358)	(0.378)	(0.346)	(0.278)	(0.432)	(0.294)
				Panel III: Matched sample, controlling for changes in housing prices					
(Δ imports from China)/worker	-0.096	0.930	-0.097	-0.168	1.018***	-0.157	0.039	0.700	0.063
	(0.286)	(0.578)	(0.283)	(0.330)	(0.392)	(0.310)	(0.354)	(0.500)	(0.356)
Δ housing price index	9.264***	10.089***	9.050***	8.559***	6.647***	8.550***	11.524***	10.963***	11.437***
	(0.825)	(1.829)	(0.741)	(0.997)	(1.254)	(0.896)	(0.934)	(1.522)	(0.831)

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
All regressions include the full set of control variables from Table 2. All regressions are weighted by start of period commuting zone's share of national population.

on wages become insignificant, while wages for college workers in the manufacturing sector even rise significantly due to the China shock.

5 Conclusion

The rapid growth in imports from China has been held responsible for the great US employment sag by several studies (ADH, 2013, Acemoglu et al, 2016, Pierce and Schott, 2016). In this note, we re-examine the empirical results by ADH (2013), who have shown in their influential work that rising exposure to imports from China led to a reduction of about 1.5 million manufacturing jobs, with a further 0.5 million jobs lost outside the manufacturing sector,[9] or roughly 2 million lost jobs in total.

Different results, however, are obtained when we also take into account the concurrent impact of the housing boom. The responses of total employment, unemployment, or not-in-the-labor force to import exposure fall by about one-half as compared to ADH, for college or noncollege workers, and become statistically insignificant in a number of cases. These altered results are obtained because housing prices act as an omitted variable in the regressions, and as shown in Figure 1, the change in this variable is negatively correlated with the change in China's import exposure especially over 2000-2007. In other words, regions of the country that experienced the greatest

[9]See note 12 for the calculation of jobs lost outside of the manufacturing sector.

increase in housing prices, like in California and Florida, had smaller import exposure from China as compared to other regions, like in the Midwest. The housing construction accompanying the rise in prices means that the *difference* between employment changes in nonmanufacturing employment between these regions was *amplified*, thereby leading to an upward bias (in absolute value) of the total employment or unemployment coefficients in our regressions.

Controlling for housing prices corrects that upward bias arising from the 'masking' effect of the housing boom, and we should expect differing results from that correction in various sectors. We have found that noncollege workers in the manufacturing sector continue to experience a reduction in employment, but that reduction does not occur in the nonmanufacturing sector. For college workers, their employment in the nonmanufacturing sector even rises with import exposure from China when controlling for the housing boom. In the 2000-2007 period, that rise in employment fully offsets their reduced employment in manufacturing, so there is no statistically significant change in their overall employment due to the China shock. Taking into account both the reduction in manufacturing employment and the increase in nonmanufacturing employment, we find that the net reduction in total US employment was about 0.8 million workers, or less than one-half of that implied by the estimates in ADH (2013).

Beyond the difference in the magnitude of job losses between our estimates and theirs, the message of our paper is that the China shock caused offsetting job gains in the nonmanufacturing sector for college educated workers. This is a surprising finding because we have not calculated the employment gains due to increased US *exports* to China, but rather, we continue to work with the exposure to Chinese *imports* as proposed by ADH (2013). Of course, for general equilibrium reasons, resources that are freed up due to import competition can be expected to be re-employed, with some lag, into export or domestic activities. The data from ADH (2013) suggest that some employment opportunities in the nonmanufacturing sector were created by the China shock, and exploring this topic further is an important direction for further research.

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Appendix

I. Calculating employment changes due to import exposure

Using the CensusACS data, ADH calculate that the US mainland population was 157.6, 178.7, and 194.3 million adults ages 16 through 64 in 1990, 2000, and 2007 respectively. Therefore, they find a supply-shock driven net reduction in US manufacturing employment of approximately 1.53 million workers:

- $[0.5\times(157.6+178.7)\times1.14+0.5\times(178.7+194.3)\times1.84]\times(0.00596\times0.48)=1.53$

Note that 0.00596 is the coefficient from the benchmark regression on manufacturing employment share, while 0.48 is the proportion of the variation in rising Chinese import exposure that can be attributed to the supply-driven components (see ADH, 2013, p. 2140 for details). We adopt this 0.48 proportion in our quantitative exercise.

We do not have detailed information on employment by education (i.e., employment for college workers versus noncollege workers), so we use available information from ADH to back out this number:

- First, the working-age population by commuting zone can be extracted from Acemoglu et al (2016).[10] Using the share of a CZ's population with a college education, available from ADH in the years 1990 and 2000, we can estimate for each commuting zone the working-age population with or without a college education in 1990 and 2000.[11]

- Using the percentage change of working-age population with a college education from 2000 to 2007, we can calculate the working-age population with college education in 2007.

- Aggregating the commuting zone level working-age population by education to the national level, we obtain the shares of college-educated workers for 1990, 2000, and 2007 as 48.2%, 53.6%, and 57.3% respectively.

- Applying these shares to the total working-age population reported in ADH (2013, note 31), we obtain total working-age population with college education as 75.9, 95.8, and 111.3 million in 1990, 2000, and 2007 respectively. Similarly, for noncollege workers the numbers are 81.7, 82.9, and 83.0 million in 1990, 2000, and 2007 respectively.

- From the benchmark results of this paper (2017, Table 2, Panel III), the percentage change of employment share due to import exposure is -0.514 percentage points for college manufacturing workers, 0.180 for college nonmanufacturing workers, -0.519 for noncollege manufacturing workers, and -0.056 for noncollege nonmanufacturing workers.

- Applying the same approach as in ADH (2013), we can calculate the employment impact of import exposure, for college and noncollege workers separately as:

[10] Note that the national working age population aggregated from this data is 163, 186, and 201 million, in 1990, 2000, and 2007 respectively, which is slightly higher than ADH's number in their note 31.

[11] Since we use the population share of college-educated person, we may underestimate the actual share of college-educated workers relative to total *working-age* population.

a. $[0.5\times(75.9+95.8)\times1.14+0.5\times(95.8+111.3)\times1.84]\times(-0.00514\times0.48)=-0.712$ million for college, manufacturing jobs;

b. $[0.5\times(75.9+95.8)\times1.14+0.5\times(95.8+111.3)\times1.84]\times(0.00407\times0.48)=0.563$ million for college, nonmanufacturing jobs;

c. $[0.5\times(81.7+82.9)\times1.14+0.5\times(82.9+83.0)\times1.84]\times(-0.00519\times0.48)=-0.614$ million for noncollege, manufacturing jobs;

d. $[0.5\times(81.7+82.9)\times1.14+0.5\times(82.9+83.0)\times1.84]\times(-0.00056\times0.48)=-0.066$ million for noncollege, nonmanufacturing jobs.

Putting together these estimates for the full period 1990-2007, Chinese import exposure caused a reduction in manufacturing employment of 1.33 million workers, while it also caused an *increase* in nonmanufacturing employment of 0.497 million workers. Thus the net reduction in total US employment due to rising Chinese import exposure is about 0.83 million workers, or less than one-half of the job less implied by the estimates in ADH (2013).[12] This striking contrast comes mainly from the different response from the nonmanufacturing sector. In particular, as shown in Table 2 and in the calculations above, rising Chinese import exposure led to a negligible reduction in noncollege nonmanufacturing jobs (0.066 million), while it also increased employment for college nonmanufacturing workers by 0.56 million.

II. Housing Prices

We use two sets of housing price data in this paper. Both can be downloaded from the Federal Housing Finance Agency website: https://www.fhfa.gov/DataTools/Downloads/pages/house-price-index.aspx.

a. Tables 2, 4, 5, and 6 use decadal changes in the annual housing price index, constructed using the annual Housing Price Index at U.S. county level. After aggregating and matching country to commuting zones, we have housing price information for 522 commuting zones in 1990, 2000, and 2007. These 522 commuting zones account for over 97 percent of US mainland population as of 1990. The excluded commuting zones are mainly rural regions without housing price information.

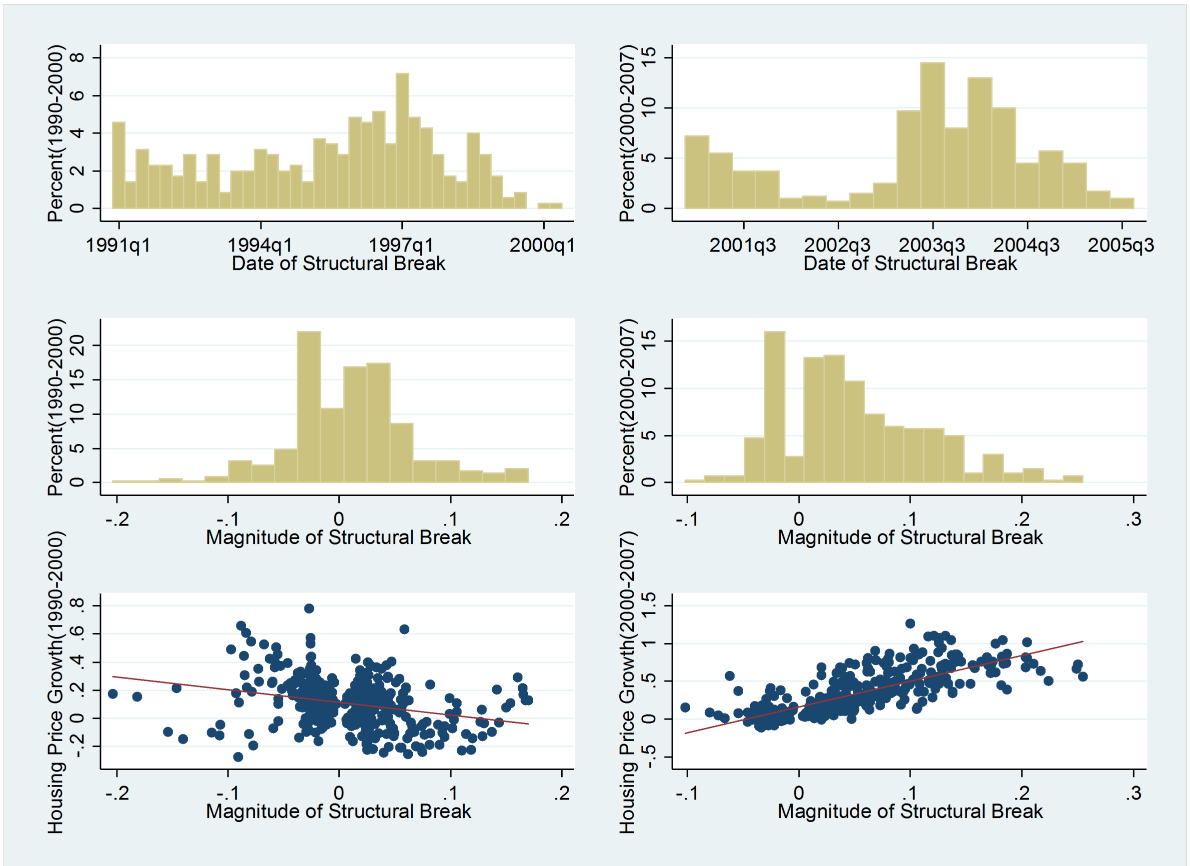
b. Table 3 and Appendix Tables A.1 and A.2 uses the housing price index at the level of metropolitan statistical area (MSA). We choose the MSA level series because the quarterly housing price information is only available at this level, which is necessary for constructing our instrument for housing price changes. Matching MSA to commuting zones lead to a further reduced sample of 321 commuting zones during 2000-2007 and 291 commuting zones during 1990-2000.

[12] Using ADH's coefficient for nonmanufacturing employment (-0.178 from Table 2, Panel I, first row), rising Chinese import exposure caused a 0.457 million job loss in the nonmanufacturing sector. Thus the reduction in total employment is 1.53+0.46=1.99, or roughly 2 million jobs. We expect that ADH (2013) did not focus on the additional job loss outside of manufacturing because that coefficient (-0.178) is highly insignificant. But when college and noncollege students are separated, then the coefficient on noncollege students outside of manufacturing (-0.531, from Table 2, Panel I, third row of estimates) is highly significant. Also, the positive coefficient on college students outside of manufacturing (0.407, from Table 2, Panel III, third row of estimates) is highly significant when housing prices are controlled for. So we consistently include both the manufacturing and the nonmanufacturing sectors in all our calculations of the overall employment changes.

III. Extra Results for Table 3

We plot the estimated dates and sizes of the structural breaks in Figure A.1. The top panel shows the dates of the breaks for each MSA, indicating that the structural breaks actually occurred at different times for different MSAs in both periods, consistent with Ferreira and Gyourko (2011) and Charles et al. (2016b). The middle panel shows the distribution of the breaks' magnitudes: during the first period (the middle-left graph), nearly half of the breaks are negative, indicating a slowdown, while most of the breaks in the second period (the middle-right graph) are postive. This relationship can also be seen in the bottom panel, during the first period (the bottom-left graph) there is a negative correlation between housing price changes and the size of the breaks, while housing price changes in the second period (the bottom-right graph) are clearly positively correlated with the size of the structural breaks.

Figure A.1: Structural Break Estimates and Relationship with Housing Price Change



In Figure A.2, we provide a few city examples for the housing price index and the structural break. It shows the quarterly data on housing price for two MSAs (Portland, OR, and Reno, NV) over the two time periods. The dashed lines reports the log of housing price while the solid line reports the structural break estimation. We use a vertical line to indicate the estimated date of the structural break. The top panel illustrates the case of Portland and the bottom panel is Reno, both of which experienced significant housing price slowdown in the first period while an obvious housing boom in the second period.

Figure A.2: Examples of Structural Breaks for Different MSAs

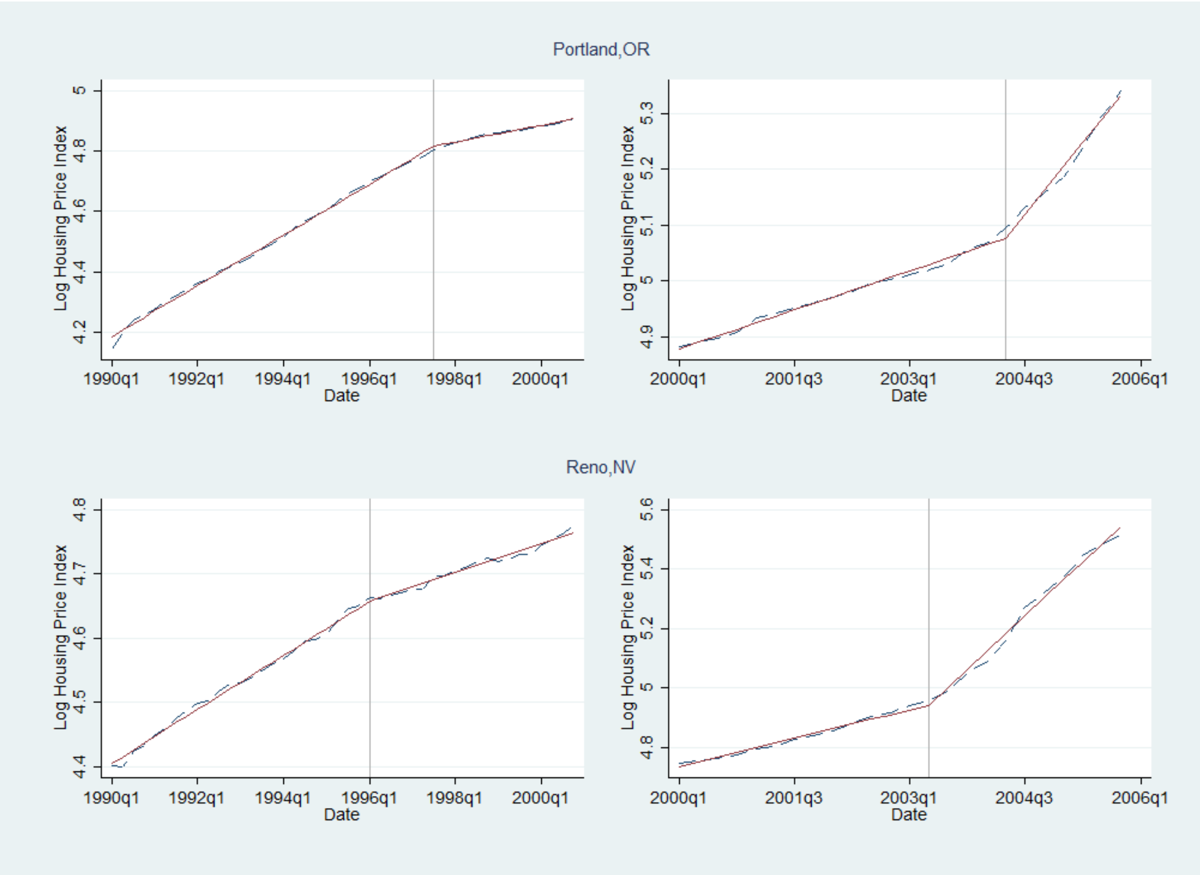


Table A.1 reports the the regressions using the ADH specification, but with the same MSA sample as in Table 3. Table A.2 controls for housing price changes, but without instrumenting for housing price changes.

Table A.1: MSA results without housing price: 1990-2007

	(1) Mfg emp	(2) Non-Mfg emp	(3) Total emp	(4) Unemp	(5) NILF
	Panel I: All education levels				
(Δ imports from China)/worker	-0.704***	-0.212	-0.916***	0.256***	0.660***
	(0.103)	(0.205)	(0.240)	(0.071)	(0.217)
	Panel II: College education				
(Δ imports from China)/worker	-0.713***	0.229	-0.484***	0.155***	0.329**
	(0.146)	(0.160)	(0.169)	(0.046)	(0.155)
	Panel III: No college education				
(Δ imports from China)/worker	-0.678***	-0.645**	-1.323***	0.305***	1.018***
	(0.106)	(0.301)	(0.351)	(0.113)	(0.312)
Observations	612	612	612	612	612
R^2	0.299	0.402	0.261	0.416	0.422

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
The MSA sample includes 321 commuting zones during 2000-2007 and 291 commuting zones during 1990-2000 that we have matched with housing prices data and the instruments for housing price changes, both of which are at the level of the metropolitan statistical area. All regressions include the full set of control variables from Table 2. All regressions are weighted by start of period commuting zone's share of national population.

Table A.2: MSA results with housing price: 1990-2007

	(1) Mfg emp	(2) Non-Mfg emp	(3) Total emp	(4) Unemp	(5) NILF
	Panel I: All education levels				
(Δ imports from China)/worker	-0.567***	0.163	-0.404	0.131*	0.273
	(0.093)	(0.247)	(0.271)	(0.072)	(0.253)
Δ housing price index	1.936***	5.324***	7.259***	-1.774***	-5.486***
	(0.497)	(0.935)	(1.349)	(0.520)	(1.046)
	Panel II: College education				
(Δ imports from China)/worker	-0.567***	0.432***	-0.135	0.065	0.071
	(0.136)	(0.160)	(0.162)	(0.051)	(0.151)
Δ housing price index	2.064***	2.881***	4.945***	-1.280***	-3.665***
	(0.532)	(0.278)	(0.516)	(0.380)	(0.402)
	Panel III: No college education				
(Δ imports from China)/worker	-0.547***	-0.097	-0.644	0.142	0.502
	(0.107)	(0.357)	(0.413)	(0.114)	(0.374)
Δ housing price index	1.849***	7.768***	9.616***	-2.315***	-7.302***
	(0.523)	(1.619)	(2.066)	(0.710)	(1.572)

Note: Robust standard errors in parentheses, clustered on state. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
see note with Table A.1.