Response to Robert Feenstra, Hong Ma, and Yuan Xu's Comment on Autor, Dorn, and Hanson (*AER* 2013)

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A 2017 comment by Feenstra, Ma, and Xu (FMX) claims that the estimation results on the impact of import competition on labor-market outcomes in Autor, Dorn and Hanson's 2013 *AER* article (ADH) are biased by the exclusion of controls for contemporaneous changes in housing prices. In investigating these concerns, we find that (a) the trend component of local housing-price changes on which FMX primarily rely is highly likely to be endogenous to trade shocks, and (b) when one incorporates measures of plausibly exogenous changes in housing prices into the main ADH specifications, there is little change in ADH's main empirical results. The key conclusions in FMX's comment stem from either using endogenous house price changes—which are themselves affected by the China trade shock, as documented by extant literature—or by employing an instrumentation strategy that inadvertently uses the China trade shock as an instrument for house price changes and then incorrectly infers that house price changes (not the China shock) are the cause of adverse employment outcomes. We conclude that the main critique provided by the comment is not supported empirically.

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In a recent comment, Robert Feenstra, Hong Ma, and Yuan Xu (2017, FMX hereafter) report that the estimation results in Autor, Dorn and Hanson (*AER* 2013, ADH hereafter) on the impact of import competition on labor-market outcomes are sensitive to the inclusion of controls for changes in housing prices. Based on their analysis, FMX conclude that ADH partially confound trade shocks with housing-market shocks. This short note contains our response to their comment.

To recap, ADH analyze the impact of import competition from China on local labor markets in the U.S. They use the 722 commuting zones (CZs) in the continental U.S. to represent local labor markets and examine the impact of the China trade shock over two time periods, 1990 to 2000 and 2000 to 2007. They measure the trade shock using the growth in China imports per worker (see their equation 3) and instrument for this value using growth in imports from China in other high-income countries (see their equation 4). ADH estimate the impact of the China trade shock on changes in CZ manufacturing employment; the CZ working-age population; CZ non-manufacturing employment, unemployment, population not in the labor force, and population receiving Social Security Disability Insurance; CZ average log weekly wages; CZ government transfer receipts per capita; and CZ household income by source. They find that CZs more exposed to import competition from China experience significantly larger reductions in manufacturing employment, increases in unemployment and non-participation in the labor force, decreases in wages in the non-manufacturing sector, increases in government transfer receipts, and reductions in household income. They do not find significant impacts of trade shocks on changes in CZ non-manufacturing employment, working-age populations, or manufacturing wages.

FMX's main critique: The comment by FMX claims that the estimation results in ADH on the impact of import competition on labor-market outcomes are sensitive to the inclusion of controls for changes in housing prices. Based on their analysis, FMX conclude that ADH partially confound trade shocks with housing-market shocks.

To begin, it is useful to review the primary regression specification in ADH, which is

$$\Delta y_{it} = \gamma_t + \beta_1 \Delta IPW_{uit} + X'_{it}\delta + \epsilon_{it}, \tag{1}$$

where Δy_{it} is the decadal change in labor-market outcome y in commuting zone i, the change in import exposure ΔIPW_{uit} is the key explanatory variable of interest, and X_{it} is a vector of control variables for initial economic conditions and demographic characteristics for CZ i. When estimating this model for the interval between 1990 and 2007, ADH stack the 10-year equivalent first differences for two periods, 1990 to 2000 and 2000 to 2007, and include separate time dummies for each decade (in the vector γ_t). The change in import exposure ΔIPW_{uit} is instrumented by the variable ΔIPW_{oit} as described in ADH equation (4). This stacked first-difference model is a three-period fixed effects model but with less restrictive assumptions on the error term (see ADH footnote 26). The vector X_{it} contains dummies for Census geographic divisions and controls for CZs' start-of-decade share of the working-age population that is college-educated, share that is foreign born, share of women in total employment, share of employment in routine occupations, and average offshorability index (see notes to Table 3 in ADH). These controls allow changes in outcomes to be a function of initial conditions, time trends to vary by geographic region, and the aggregate time trend to vary by decade. Standard errors are clustered at the state level.

FMX claim that the estimation results in ADH on the impact of import competition on labor-market outcomes are biased by the exclusion of controls for contemporaneous changes in housing prices. Specifically, the claim is that trade shocks were most pronounced in CZs where house prices fell by more, thus confounding the impact of the China shock with the adverse impact of housing shocks. FMX draw inspiration from Charles, Hurst, and Notowidigdo (2016a and 2016b, CHN hereafter), who observe that the housing boom in the 2000s temporarily 'masked' the adverse effects of employment shocks to manufacturing. Perhaps ironically, FMX's claims are opposite in spirit to CHN; FMX's reasoning implies that adverse housing shocks *amplified rather than masked* the measured impact of trade shocks.

FMX propose adding to equation (1) the variable $\triangle P_{it}$, which is the contemporaneous percentage change in housing prices for commuting zone i, a measure which is available for 522 of the 722 commuting zones used in the ADH analysis. FMX find that some ADH coefficient estimates (on the impact of trade shocks on CZ changes in total employment, unemployment, and non-participation in the labor force) fall in magnitude and (or) lose statistical significance when one includes CZ-level housing-price changes in the regression. By failing to control for the "masking effect" of the US housing-price boom, FMX suggest that ADH overstate the labor-market impacts of the China trade shock.

Observation 1: Changes in local housing prices are fundamentally endogenous to changes in local labor demand

There is abundant empirical and theoretical literature which establishes that shocks to local labor demand affect local housing values (e.g., Gyourko and Glaeser, 2005; Notowidigdo, 2011; Diamond, 2016). Specifically in the context of import competition, Feler and Senses (2016) find that commuting zones more exposed to the ADH China trade shock (where they employ the ADH instrumentation strategy) experience larger reductions in housing prices, business activity, tax revenues and expenditure on social welfare programs and pubic housing. In conjunction with the results in ADH, these results indicate that adverse shocks to manufacturing employment, be they due to trade or other forces, are strongly likely to weaken local demand for housing and thereby produce a contraction in housing values. The comment, however, appears to give short shrift to the sensitivity of local housing prices to local-labor-market conditions. FMX do not mention the endogeneity issue until rather late in the paper (with no mention in the abstract or introduction) and instrument for housing prices in only one of the three sections of estimation results. When an instrumentation strategy for housing price changes is introduced (in section 3), it relies on assumptions that are difficult to support.

Observation 2: The instrumentation approach in FMX makes implausible identifying assumptions

The comment's proposed solution to the endogeneity of housing prices is to use an empirical model of structural breaks in housing values to construct an instrument. Specifically, FMX regress log housing prices for MSAs at a quarterly frequency on an MSA dummy, an MSA-specific time trend, and two MSA-specific breaks in the time trend (one for the 1990s and one for the 2000s). The two decadal breaks in the time trend, along with the ADH trade-shock instrument, are then jointly used to instrument for the changes in local import exposure and changes in local housing prices.

One evident problem with this approach is that it requires the identifying assumption that trend breaks in local housing prices are uncorrelated with unobserved local-labor-demand shocks, which are embodied in the disturbance ϵ_{it} in equation (1). The literature cited above strongly suggests that this assumption is invalid. Indeed, available empirical literature implies that changes in local labor demand are a driving force behind observed changes in housing values. Just as real-estate prices have fallen in Youngstown, Ohio, with the shuttering of local manufacturing plants, they have soared in the San Francisco Bay Area with the proliferation of high-tech startups. The FMX instrumental variables strategy requires that *deviations from a linear change in house prices are exogenous*—which is entirely a functional form restriction on the smoothness of underlying price shocks. In reality, it is hard to think of any economic shock that *necessarily* causes a linear but not a nonlinear change in prices. FMX's IV approach may be sufficient for identifying discreet trend breaks in house price trends—which is the purpose for which Charles, Hurst and Notowidigdo developed it—but it is silent on the underlying causes of these shocks. It is almost certain that the China trade shock *is* one of those causes.

There is a further conceptual error in the FMX instrumentation strategy that renders the results largely uninformative—and in fact, which causes the authors to misinterpret their own results. Because both sets of instruments—the ADH instrument based on import growth in other high-income countries and the estimated structural breaks in local housing prices—are used jointly to instrument for two endogenous variables (the ADH trade shock and the local housing-rice changes), part of the predicted change in local housing prices is due to the ADH trade-shock instrument. This is directly shown in column 2 of the lower panel of Table 3 of the FMX comment, where the ADH instrument is a *significant, negative predictor* of local housing prices changes even conditional on FMX's housing instruments. Thus, under the FMX approach, exogenous trade shocks are allowed (statistically) to affect local-labor-market outcomes partly through the variable ΔIPW_{uit} and partly through the variable $\triangle P_{it}$. In this statistical model, a change in the coefficient value on ΔIPW_{uit} in equation (1) when including the variable $\triangle P_{it}$ in the regression would not mean that the impact of the trade on the outcome variable Δy_{it} is lessened; rather, it would mean that trade shocks affect labor-market outcomes through multiple channels. Specifically, the FMX instrumental variables strategy measures the causal effect of the China shock on labor market outcomes that accrues both through shocks to manufacturing *and* through shocks to house prices. Though FMX don't do the covariance calculations that would be needed to scale this *joint* effect, it's plausible that the total effect of the China shock that works through both the manufacturing and housing channels is as large or larger than the marginal effect operating purely through manufacturing shocks. (Simply including the trade shock as a reduced-form, non-instrumented covariate would account for both channels at once without having to specify multiple endogenous variables).

Observation 3: When one uses an alternative measure of housing-price changes that is plausibly exogenous, the FMX critique is largely neutralized

As an alternative to the FMX approach, we employ a strategy to capture local housing-price changes that cannot be readily explained by national trade shocks. Using the FMX sample of 522 commuting zones and their data on house prices, we regress the relative change in CZ house prices on ΔIPW_{uit} (or ΔIPW_{uit} plus full ADH controls) in all *out-of-state* CZs. We then use the estimates to compute residual house price changes for *in-state* CZs.[1] The resulting residualized housing-price changes are highly correlated with the the gross values that FMX use (with a correlation coefficient between 0.72 and 0.98, depending on whether no controls or full controls are used). Table 1 reports our replication of the FMX results on their restricted sample of 522 commuting zones, where we alternatively include the ADH trade shock alone (panel I), the ADH trade shock plus the FMX (plausibly *endogenous*) housing-price measure (panel II), or the ADH trade shock plus (plausibly *exogenous*) residualized housing-price changes (panel III).[2] We report each specification for three outcome measures—manufacturing employment, non-manufacturing employment, and total employment—either using no additional control variables other than a period dummy (panel A) or full ADH controls (panel B). Table 2 then repeats the analysis separately for the two time periods, 1990-2000 and 2000-2007.

Beginning with Table 1 for the stacked time period, we see that ADH results are largely unaffected when we go from the ADH specification (panel I) to the specification that includes controls for changes in housing prices that are plausibly *exogenous* to the import shock (panel III). In all specifications, the impact of the ADH trade shock on manufacturing employment is strongly negative

[1]The exogenous component of the CZ-level house price shock is calculated as follows: (1) regressing the change in the house price index in each CZ on the instrumented China trade shock using a leave-one-out specification where data from the state each encompassing each CZ is omitted (a total of 48 regressions); (2) predicting the China-shock-induced (endogenous) component of CZ-level house price changes by applying the coefficients from the out-of-state regression to CZ-specific values, including the instrumented CZ-specific trade shock; (3) calculating the exogenous CZ-level house price shock as the difference between the observed house price change and the predicted (endogenous) change based on the out-of-state regression coefficient. The first stage regression of house price changes on the instrumented China trade shock from the leave-one-out specification has a mean coefficient of -0.046 and a mean t-ratio of -5.36. When the 14 ADH standard covariates are added to the model, the mean coefficient increases in absolute magnitude to -0.104 with a mean t-ratio of -3.74. The coefficients and standard errors from these 48 regressions are remarkably stable across each regression; the standard deviations of the coefficients and standard-errors are 0.0037 and 0.0003, respectively, for the base specification and 0.005 and 0.009 for the augmented specification. This stability indicates that no one state exerts substantial influence on the overall pattern.

[2]The results in panel II of Table 1 differ slightly from those in panel III of Table 2 in FMX because we compute decadal changes for both ΔIPW_{uit} and $\triangle P_{it}$, whereas FMX combine a decadal change in ΔIPW_{uit} with a seven-year change in $\triangle P_{it}$ in the second period.

and precisely estimated, with little change in magnitude across panels. If we compare coefficient estimates on the ADH trade shock for total employment (column 3), which FMX claim is sensitive to changes in housing prices, we see little change in magnitude (and continued high precision), whether or not we include residualized housing-price changes (panel I versus panel III). As in ADH, the impact of trade shocks on non-manufacturing employment is generally imprecisely estimated (with the addition of housing-price controls not affecting this finding). Only in Panel II, which mixes endogenous and exogenous variation in house price changes as per FMX, do we see a substantial reduction in estimated effect of the China Shock on total employment. As explained above, this approach is not conceptually sound.

Table 2 analyzes the 1990s and 2000s separately with models that include exogenous housing prices and either comprise no covariates or full covariates. As in ADH, we see that in the 1990s, when the China shock was relatively weak and less diversified across industries, we obtain less precisely estimated coefficients than in the 2000s. This holds whether we use the ADH trade shock, measured as imports per worker (panel I), or the alternative trade-shock measure of the change in import penetration (as in Acemoglu, Autor, Dorn, Hanson, and Price, 2016). The impacts of trade shocks on manufacturing employment and total employment are robustly negative and very precisely estimated (for all but the case of import penetration with no controls in column 3, panel A.II, where the trade shock is marginally significant). As in ADH, the impacts of trade shocks on non-manufacturing employment are imprecisely estimated when full controls are included in the specification.

To facilitate comparison, columns (4) and (8) of the table report the ratio of the estimated impact of the China trade shock on non-manufacturing vs. manufacturing employment, both as a share of working-age population. The ratio of these coefficients is typically around -0.5 to -0.6, implying that employment gains in non-manufacturing offset about 50 to 60 percent of the losses in manufacturing (i.e., less than full offset). The estimates for the 2000s are more sensitive to the inclusion of controls than are those for the 1990s. But when the *exogenous* housing shock component is included with the 14 covariates used by ADH (lower row of each panel), the pattern of incomplete employment offset in non-manufacturing is evident.

Observation 4: FMX inadvertently mischaracterize a core finding of ADH

In contrast to what FMX claim in their conclusion, ADH do *not* compute job loss outside of manufacturing (precisely because the results for non-manufacturing employment are mixed). Nevertheless, throughout the ADH analysis there is no evidence that CZ employment gains outside of manufacturing fully offset CZ employment losses within manufacturing. The question whether workers are able to smoothly transition from the declining manufacturing sector to other equally lucrative employment outside of manufacturing is of course best addresses by tracing out the individual career paths of workers in longitudinal data. In a separate research paper, Autor, Dorn, Hanson, and Song (2014) analyze individual-level longitudinal data from the Social Security Administration and find that over a near two-decade period, trade-exposed workers do not fully replace earnings losses in trade-exposed industries with earnings gains in other industries. The implication of FMX's analysis that displaced manufacturing workers simply transition to other equally paid jobs after the China shock is not supported by this evidence.

Summary

The comment by FMX claims that the estimation results in ADH on the impact of import competition on labor-market outcomes are biased by the exclusion of controls for contemporaneous changes in housing prices. However, further investigation reveals that (a) the trend component of local housing-price changes is highly likely to be endogenous to trade shocks, and (b) when one incorporates measures of plausibly exogenous changes in housing prices into the main ADH specifications, there is little change in ADH's main empirical results. The key findings in FMX's comment stem from either using endogenous house price changes—which are themselves affected by the China trade shock, as documented by extant literature—or by employing an instrumentation strategy that inadvertently uses the China trade shock as an instrument for house price changes and then incorrectly infers that house price changes (not the China shock) are the cause of adverse employment outcomes. We conclude that the main critique provided by the comment is not well-founded methodologically or empirically.

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Table 1. Imports from China and Employment Status of Working Age Population within Commuting Zones, 1990-2007: 2SLS Estimates.

	A. Models w/o Additional Control Variables						B. Models with Full ADH Control Variables					
	Mfg Emp (1)		Non-Mfg Emp (2)		Total Emp (3)		Mfg Emp (4)		Non-Mfg Emp (5)		Total Emp (6)	
	I. No Control for House Price Change											
(Δ Imports from China to US)/Worker	-0.791	**	0.103		-0.688	**	-0.661	**	-0.177		-0.837	**
	(0.060)		(0.101)		(0.109)		(0.100)		(0.161)		(0.198)	
	II. Control for Gross (Exogenous + Endogenous) House Price Change											
(Δ Imports from China to US)/Worker	-0.707	**	0.381	**	-0.326	**	-0.522	**	0.197		-0.325	
	(0.063)		(0.097)		(0.110)		(0.079)		(0.207)		(0.219)	
Relative Change in Gross House Price Index	0.922	**	3.056	**	3.978	**	1.333	**	3.590	**	4.923	**
	(0.290)		(0.444)		(0.701)		(0.292)		(0.543)		(0.777)	
	III. Control for Exogenous House Price Change											
(Δ Imports from China to US)/Worker	-0.746	**	0.225	*	-0.521	**	-0.655	**	-0.166		-0.822	**
	(0.061)		(0.094)		(0.103)		(0.072)		(0.177)		(0.182)	
Relative Change in HPI, Exogeous to Import Shock	1.037	**	2.794	**	3.831	**	1.583	**	3.148	**	4.731	**
	(0.234)		(0.414)		(0.606)		(0.215)		(0.424)		(0.559)	

Notes: N=1044 (522 commuting zones x 2 time periods). Regressions in Panel A control for a period dummy and for the change in the house price index (HPI) in subpanels A.II and A.III. Regressions in Panel B control for an additional 14 covariates as indicated in column 6 of Table 3 in Autor, Dorn and Hanson (2013). The exogenous change in HPI in panel A.III for a given CZ is computed as the residual of a regression of the change in out-of-state HPI on imports-per-worker and a period dummy (i.e., change in HPI is first regressed on IPW for all local labor markets outside the CZ's own state, and the predicted value of that regression for the CZ is then subtracted from the gross change in HPI for that CZ). The exogenous change in HPI in panel B.III is computed accordingly while controlling for the additional covariates. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The exogenous component of the CZ-level house price shock is calculated by: (1) regressing the change in the house price index in each CZ on the instrumented China trade shock using a leave-one-out specification where data from the state each encompassing each CZ is omitted (a total of 48 regressions); (2) predicting the China-shock-induced (endogenous) component of CZ-level house price changes by applying the coefficients from the out-of-state regression to CZ-specific values, including the instrumented CZ-specific trade shock; (3) calculating the exogenous CZ-level house price shock as the difference between the observed house price change and the predicted (endogenous) change based on the out-of-state regression coefficient. The first stage regression of house price changes on the instrumented China trade shock from the leave-one-out specification has a mean coefficient of -0.046 and a mean t-ratio of -5.36. When the 14 ADH standard covariates are added to the model, the mean coefficient increases in absolute magnitude to -0.104 with a mean t-ratio of -3.74. The coefficients and standard errors from these 48 regressions are remarkably stable across each regression; the standard deviations of the coefficients and standard-errors are 0.0037 and 0.0003, respectively, for the base specification and 0.005 and 0.009 for the augmented specification. This stability indicates that no one state exerts substantial influence on the overall pattern.

Table 2. Imports from China and Employment Status of Working Age Population within Commuting Zones, Subperiods 1990-2000 and 2000-2007: 2SLS Estimates. Dep Vars: 10-Year Equivalent Changes in Population Shares by Employment Status

	A. Subperiod 1990-2000							B. Subperiod 2000-2007						
	Mfg Emp		Non-Mfg Emp		Total Emp		β(2)/β(1)	Mfg Emp		Non-Mfg Emp	Total Emp		β(2)/β(1)
	(1)		(2)		(3)		(4)	(5)		(6)	(7)		(8)
							I. Coefficient Estimates for CZ Imports per Worker						
Model w/o control variables	-0.990	**	0.495	~	-0.495	**	-0.5	-0.755	**	0.032	-0.723	**	0.0
	(0.197)		(0.284)		(0.109)			(0.059)		(0.134)	(0.148)		
Model w/ full controls + exogenous HPI	-0.320		0.180		-0.140		-0.6	-0.671	**	0.252	-0.419	*	-0.4
	(0.234)		(0.269)		(0.252)			(0.148)		(0.160)	(0.179)		
							II. Coefficient Estimates for CZ Import Penetration						
Model w/o control variables	-2.138	**	1.228	**	-0.910	~	-0.6	-2.077	**	0.142	-1.936	**	-0.1
	(0.363)		(0.465)		(0.505)			(0.412)		(0.251)	(0.379)		
Model w/ full controls + exogenous HPI	-1.552	*	1.091	~	-0.462		-0.7	-1.060	*	0.491	-0.569	~	-0.5
	(0.668)		(0.633)		(0.900)			(0.431)		(0.313)	(0.325)		

Notes: N=1044 (522 commuting zones x 2 time periods). Each coefficient comes from a separate regression of the outcome on a measure of Chinese import competition (and additional controls in rows 2 and 4). Panel I measures import competition from China using the import-per-worker measure of Autor, Dorn and Hanson (2013). Panel II measures import competition using the import penetration measure of Autor, Dorn and Hanson (2017). Models with full controls include the full vector of control variables from panel B.III of Table 1. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.