The Causal Effect of the Dollar on Trade

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Abstract

This paper establishes a causal link between the dollar exchange rate and international trade flows, employing a new instrument for the U.S. Dollar that is based on domestic U.S. housing activity (Ma and Zhang (2019)). In line with the dominant currency paradigm (Gopinath et al. (2020)), import prices and quantities respond strongly to a country's exchange rate with the U.S. dollar. Once we instrument the dollar, we find evidence for perfect pass-through of the dollar exchange rate to import prices. A dollar appreciation of 1 percent lowers import quantities by 1.5 percent for countries that fully invoice in dollars.

Keywords: Dominant Currency, Dollar Invoicing, International Trade

JEL Classification: F14, F31, G15

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1 Introduction

In an influential paper, Gopinath et al. (2020) document the dominant role of the U.S. dollar in international trade and show that if export prices are sticky and invoiced in a dominant currency, this has important implications for price pass-through and the trade elasticity.[1] Predictions under the dominant currency paradigm (DCP) differ strikingly from those obtained under its alternatives, producer currency pricing (PCP) and local currency pricing (LCP). Most importantly, the DCP predicts that changes to a country's exchange rate against the dollar should be fully passed through to import prices and should have a sizable effect on import quantities. In contrast, controlling for the exchange rate vis-a-vis the dominant currency, the bilateral exchange rate with the trading partner should not affect import prices or quantities.

Gopinath et al. (2020) provide extensive evidence for the key predictions of their model, using both aggregate data at the country-pair year level, as well as looking at detailed trade data from Colombia. A key challenge in testing the predictions from the DCP is the endogeneity of the U.S. dollar. As macroeconomic variables, trade flows, and the dollar are co-determined, it is difficult to provide evidence in favor of the model that goes beyond correlations.

This paper employs a new variable to instrument the dollar, building on recent research by Ma and Zhang (2019), who show that domestic housing activity in the United States can predict the dollar one year ahead. U.S. housing represents a particularly well-suited

[1] Earlier work on the role of the dollar includes Corsetti and Pesenti (2005), Cook and Devereux (2006), Devereux et al. (2007), Goldberg and Tille (2008), Goldberg and Tille (2009), Canzoneri et al. (2013), and Gopinath (2015). The earlier literature, in particular the influential paper by Goldberg and Tille (2008), discussed dollar invoicing by third countries as vehicle currency trade.

instrument for the question we are studying, as we are focusing our analysis on trade between third countries, whose trade flows are unlikely to be directly affected by changes to the domestic demand and supply for U.S. housing.

When instrumenting the U.S. dollar with a one year lag of new housing permits in the United States, the estimated relationships between the dollar exchange rate, import prices, and import quantities strengthen substantially relative to the OLS results in Gopinath et al. (2020). In particular, we find highly significant effects for the interaction between the dollar invoicing share of imports and the dollar exchange rate, both for import prices and import quantities. Moreover, the estimated interaction term coefficients are much larger when using our 2SLS approach. For import prices, the coefficient is very close to one (the value predicted by DCP), whereas the interaction coefficient for import quantities is larger than one (also as predicted by DCP, when assuming a trade elasticity larger than one). In a more recent paper, Boz et al. (2020) adopts the same empirical strategy of Gopinath et al. (2020) using a new data set for invoicing currency shares. Using a more refined currency share data improves the statistical significance of the estimated interaction term coefficients, which is consistent with the highly significant findings in this paper. However, different from our paper, their estimated interaction coefficient for import prices is remains significantly smaller than one, indicating an imperfect pass-through of the dollar to import prices.

More generally, the literature on dominant currency pricing and its implications has been expanding rapidly. See, for example, Amiti et al. (2020) on the invoicing choices of Belgian firms, as well as Corsetti et al. (2020) on the implications of a dominant currency for the transmission of shocks across borders and optimal monetary policy. See also Goldberg and Tille (2016) who study vehicle currency invoicing with Canadian transaction-level import data.

The paper proceeds as follows. Section 2 presents the theoretical framework. Section 3 discusses the instrumental variable approach. Section 4 explains the data sets. Section 5 shows our main results and Section 6 concludes.

2 Conceptional Framework

Below, we outline the key elements of the model in Gopinath et al. (2020) that are relevant for our empirical analysis

Baseline model and predictions Suppose there are three countries: the United States, country i, and country j. Denote the share of currency k in the invoicing of exports from country i to j as θ_{ij}^{k}. Let $e_{ij,t}$ denote the price of currency i in units of currency j and $e_{\$j,t}$ the price of the dollar in units of currency j. That is, an increase in $e_{\$j,t}$ reflects an appreciation of the dollar. Assume that firms set prices one period in advance and that they face a demand function of the form $q = p^{-\sigma}A$, where σ is the elasticity of substitutions across varieties and A is the level of aggregate demand. In addition, assume that firms' marginal costs are constant in local currency. It is then easy to show that the following relationship holds (log-linear approximation):

Proposition 1 (Prices and Quantities) *Suppose prices are rigid one-period ahead. Then* *pass-through into prices and quantities is given by:*

$$\Delta p_{ij,t} = \theta^{i}_{ij}\Delta e_{ij,t} + \theta^{\$}_{ij}\Delta e_{\$j,t} \tag{1}$$

$$\Delta y_{ij,t} = -\sigma\left(\theta^{i}_{ij}\Delta e_{ij,t} + \theta^{\$}_{ij}\Delta e_{\$j,t}\right), \tag{2}$$

where θ^{i}_{ij} is the share of currency i in imports of country i from country j.

As pointed out in Gopinath et al. (2020), it is helpful to look at the three corner cases for currency invoicing. First, consider *local currency pricing (LCP)*, that is firms price their goods only in the destination country currency, $\theta^{j}_{ij} = 1$. Then, $\Delta p_{ij,t} = \Delta y_{ij,t} = 0$ and there is no pass-through into export prices and quantities. In contrast, under *producer currency pricing (PCP)*, when $\theta^{i}_{ij} = 1$, trade quantities and prices should only react to the bilateral exchange rate between i and j. Finally, if all firms follow dominant currency pricing with $\theta^{\$}_{ij} = 1$, then only the exchange rate between destination country j and the dollar, $e_{\$j}$, should matter for pass-through. Additionally, the effect of the dollar exchange rate should be the larger, the larger is the share of imports that is invoiced in dollars.

3 Empirical Approach

The key challenge in testing equations (1) and (2) is that the exchange rates are co-determined with other macroeconomic factors that move trade prices and quantities. To test for the causal relationship implied by the model hence requires a shock that moves the exchange rate but does not directly impact trade prices and quantities between countries i

and j other than through the exchange rate. In the following, we show that U.S. domestic housing activity represents such a shock. That variable is able to forecast moves in the dollar one-year ahead, while being plausibly exogenous to the bilateral trade between two countries other than the United States.

US housing cycles and the Dollar Ma and Zhang (2019) uncovers that US housing capital investment, such as residential investment and building permits, are strong in-sample and out-of-sample predictors for the dollar up to three years.

One plausible explanation why housing capital investment affects the future price of the dollar is through their effect on the relative supply of traded and non-traded goods, as housing is one of the most important non-traded goods. Ma and Zhang (2019) propose a model where the price of the traded good is determined globally but the domestic price of the non-traded good is mostly determined by domestic supply and demand. In that setting, output fluctuations in the domestic non-traded good can generate strong adjustments in the relative price between the non-traded and the traded good, and hence impact the value of the dollar. This is known as the relative price adjustment channel.[2]

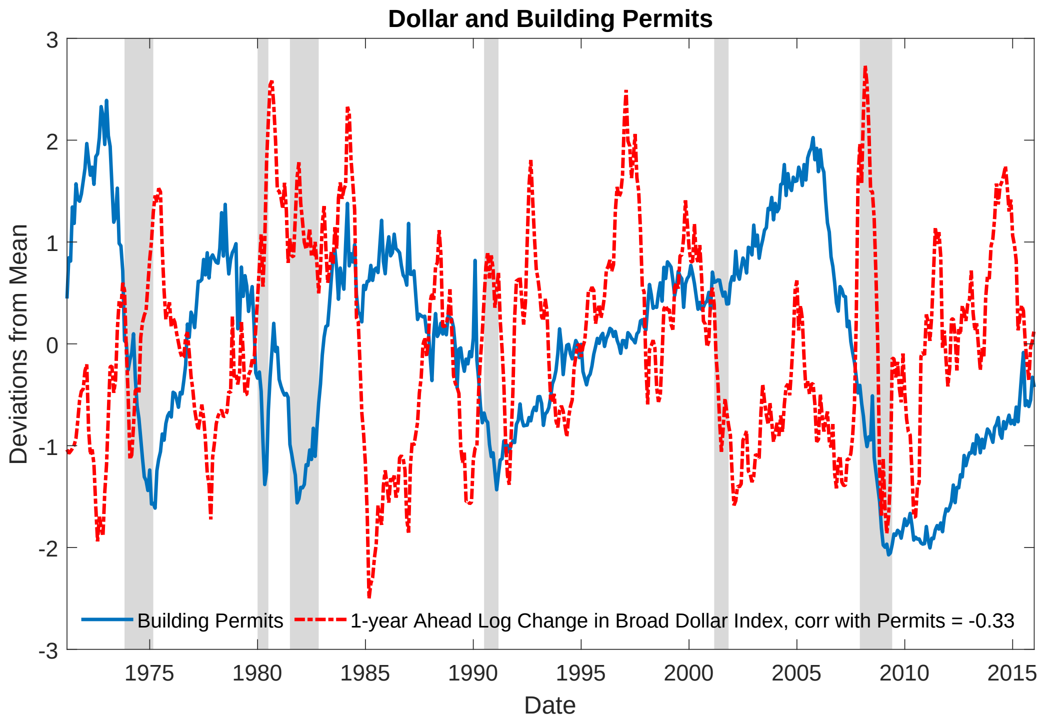
Figure 1 shows the time series of the total number of building permits authorized in the US against two-year ahead log changes in an average dollar index.[3] The two series exhibit a negative 33% correlation, indicating that higher housing capital investment predicts persistent future dollar depreciation.

While U.S. *domestic* housing capital investment is highly correlated with future dollar

[2] Ma and Zhang (2019) shows that increases in US domestic housing investment indeed predict persistent declines in the relative price of the non-traded price measure from Betts and Kehoe (2008).

[3] The dollar index in the figure is computed as an equal-weighted value of the dollar against a broad group of currencies which consists of 19 advanced economies and 13 emerging markets. See Ma and Zhang (2019) for more details

Figure 1. US Housing Investment and the Dollar



The figure plots the time-series of the standardized US building permit authorized and one-year ahead log change in average dollar index. The average dollar Index is computed as an equal-weighted average value of the U.S. dollars against a broad group of currencies which consists of 19 advanced economies and 13 emerging markets. Shaded areas correspond to NBER recession dates. The data is monthly and spans the period 1971 to 2016

movement, one would not expect domestic U.S. housing activity to directly affect bilateral trade between two other countries. While, based on the logic in the model proposed by Ma and Zhang (2019), we would expect a change in the relative price between the traded and the non-traded good in the United States to affect U.S. trade with other countries, but not directly affect trade between third countries. Of course, if relative prices between the United States and another country change, this can lead to trade diversion. We will address this concern in the estimation by including time fixed effects and by estimating the model separately for countries with high and low trade exposure to the United States.

One could be concerned, however, that international capital flows might jointly affect the

U.S. current account balance and the U.S. housing cycle, as remarked in Bernanke (2005). As a result, U.S. housing activity could be a proxy for international capital flows into the United States. However, Lilley et al. (2019) find that international capital flows were disconnected from exchange rate fluctuations in the period before 2007. Our sample starts in 1988, and the housing-dollar relationship also holds in the pre-crisis sample. In addition, Ma and Zhang (2019) find that various measures of international capital flows cannot explain the ability of housing investments to predict the dollar.

While we use US building permits as the main measure of the housing investment, we also consider two alternative measures. The first one is housing starts obtained from the Survey of Construction (SOC). Housing construction is a long process, and obtaining a building permit is the first step in this process.[4] Building permits are a measure of potential home construction starts, as not every permit leads to a construction start. In particular, building permits are temporary authorizations, and they expire if construction does not start within a certain time period. Housing starts, in contrast, directly count new home constructions that are started in a given period.

The second alternative measure is private residential fixed investment (PRFI). In contrast to permits and housing starts, which are count variables, PRFI measures investment expenditures in dollar terms. While the advantage of count measures is that they are insulated from the housing price fluctuations, they cannot capture quality improvements in real estate. To normalize the variable over time, we scale PRFI by concurrent gross domestic private investment (GDPI). Ma and Zhang (2019) show that both alternative measures, housing starts and PRFI/GDPI, can robustly predict the dollar in-sample and out-of-sample for

[4] According to 2016 SOC, the average time for the construction of a new single-family home is 6 months.

up to twelve quarters.

4 Data

Data for the bilateral trade is from Gopinath et al. (2020).[5] They construct annual panel data on bilateral trade volumes from UN Comtrade. The dataset provides detailed customs data for a large set of countries at the HS 6-digit product level with information about the destination country, dollar value, quantity, and weight of imports and exports. The analysis in the paper focuses on the non-commodity goods which are broadly defined as HS chapters 1–27 and 72–83, which comprise animal, vegetable, food, mineral, and metal products. The dollar invoicing share for each importer is from Table 10 of Gopinath et al. (2020).[6]

Data on US housing capital investment including building permits is from the Building Permits Survey (BPS) conducted by the Census. We further supplement the investment data with US private residential fixed investment (PRFI) from the national income and product accounts (NIPA) Table 1.1.5 (line 13). Housing starts are obtained from the Survey of Construction (SOC) conducted by the Census.

Our final sample includes 54 countries and over 2,800 dyads that cover more than 90% of world trade. The sample is annual and spans the periods from 1988 to 2015.

[5] We are very thankful to the authors for sharing their data with us.

[6] For robustness, we also consider the updated and time-varying dollar invoicing share data from a more recent paper Boz et al. (2020)

5 Results

Import Quantities and the Dollar We start our analysis by investigating the relationship between the dollar exchange rate and import quantities, replicating the specification in Gopinath et al. (2020). The baseline specification is as follows:

$$\begin{aligned}\Delta y_{ij,t} &= \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \beta_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \beta_k^{\$} \Delta e_{\$j,t-k} \\ &\quad + \sum_{k=0}^{2} \eta_k \Delta e_{ij,t-k} \times S_j + \sum_{k=0}^{2} \eta_k^{\$} \Delta e_{\$j,t-k} \times S_j + \theta' X_{j,t} + \varepsilon_{ij,t},\end{aligned} \tag{3}$$

where $\Delta y_{ij,t}$ is the log difference in bilateral import quantities, $e_{ij,t}$ is the log price of currency i in units of currency j, $e_{\$j,t}$ is the log price of the U.S. dollar in units of currency j, and S_j is the importing country's dollar invoicing share.

To address the endgeneity concerns discussed above, we then estimate a 2SLS specification:

$$\begin{aligned}\Delta y_{ij,t} &= \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \beta_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \beta_k^{\$} \Delta e_{\$j,t-k} \\ &\quad + \sum_{k=0}^{2} \eta_k \Delta e_{ij,t-k} \times S_j + \sum_{k=0}^{2} \eta_k^{\$} \Delta e_{\$j,t-k}^{IV} \times S_j + \theta' X_{j,t} + \varepsilon_{ij,t},\end{aligned} \tag{4}$$

where $e_{\$j,t}^{IV}$ is the value of the dollar in unit of currency j *instrumented* by the lagged log number of U.S. housing permits issued. The first-stage equation is given by,

$$e_{\$j,t} = \alpha_j + \sum_{k=1}^{P} \beta_k^{H} H_{t-k} + \epsilon_{j,t}. \tag{5}$$

where H_t is the measure of U.S. housing activity. Our main specification for estimating equation (5) uses $P = 1$ but the results are not sensitive to the choice of P, as shown in the robustness results. For all specifications, we include dyadic and time fixed effects λ_{ij} and δ_t. Following Gopinath et al. (2020), controls $X_{j,t}$ include two lags of the independent variables, and the growth rate of the real GDP of the importing country j and its two lags.

The first four columns of Table 1 reports the OLS estimates based on equation (3) while columns (5) to (7) report the 2SLS estimates based on equation (4). To conserve space, for each column, we only report the coefficients of $\Delta e_{ij,t}$ (changes in bilateral exchange rate at time t), $\Delta e_{\$j,t}$ (changes in prices of the dollar in currency j at t), as well time-t interaction terms of importing country's dollar invoicing share S with $\Delta e_{ij,t}$ and $\Delta e_{\$j,t}$, respectively.

Not surprisingly, our OLS estimates in columns (1) to (3) perfectly replicate the results in Gopinath et al. (2020). The negative interaction term between the dollar and the invoicing share S in column (3) suggest that a dollar appreciation has a more negative impact on the trade volume for importers with a larger dollar invoicing share. However, this relationship is not precisely estimated, as the coefficient is insignificant at conventional levels.

Once we instrument for the dollar in column (6), this relationship becomes highly statistically significant and the coefficient becomes much larger, with an estimated value of -1.5. Similarly, comparing results reported in column (2) and (5), we find that, once the dollar is instrumented, an increase in the value of the dollar against importer's currency implies a much larger decline in import quantities. The large change in the coefficient sizes when moving from OLS to the 2SLS suggests that the endogeneity of the dollar biases the estimated OLS coefficients strongly towards zero. Moreover, the sizes of the estimated coefficients are much more in line with what one would expect from the theoretical results on the DCP. Our

2SLS results thus not only provide causal evidence on the link between the dollar and import quantities but also provide estimates of a magnitude that further strengthen the plausibility of the DCP as studied in Gopinath et al. (2020).

A follow-up paper, Boz et al. (2020), that refines the data set on currency invoicing, finds significant interaction effects between the dollar invoicing share and the dollar exchange rate. However, the estimated interaction term coefficients in Boz et al. (2020) is only around 0.4, and thus much smaller than the coefficient of 1.5 that we find in our 2SLS estimation. This suggests that controlling for previously omitted variables like the euro invoicing share is helpful but does not fully resolve the endogeneity problem of the dollar.

For robustness, we use these updated and now time-varying dollar invoicing share data from Boz et al. (2020) in columns (4) and (7).[7] With the updated data our 2SLS estimate on the interaction term in column (7) is a bit larger than the coefficient estimated in column (6) but not statistically different.

Bilateral Pass-through Regressions Next, we investigate the link between the dollar and the international *prices*. As before, we first replicate the reduced-form exchange rate pass-through specification in Gopinath et al. (2020), and then instrument the dollar exchange rate with our housing instrument. The baseline specification now reads:

$$\begin{aligned}\Delta p_{ij,t} &= \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \beta_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \beta_k^{\$} \Delta e_{\$j,t-k} \\ &\quad + \sum_{k=0}^{2} \eta_k \Delta e_{ij,t-k} \times S_j + \sum_{k=0}^{2} \eta_k^{\$} \Delta e_{\$j,t-k}^{IV} \times S_j + \theta' X_{j,t} + \varepsilon_{ij,t}, \end{aligned} \tag{6}$$

[7] Note that results in column (4) to not perfectly match the numbers in Boz et al. (2020), as we currently do not have access to their updated trade data.

where $\Delta p_{ij,t}$ is the log difference of the price of goods exported from country i to country j measured in importer currency j. The first-stage regression remains the same as equation (5). As control variables $X_{i,t}$, we include changes in the (log) producer price index in the exporting country i and two lags of this variable, and two lags of the bilateral exchange rate.

Table 2 report both OLS and 2SLS estimates of pass-through regression (6). OLS columns (1) to (3) again perfectly replicate the results in Gopinath et al. (2020). They suggest that the import invoicing share plays an important role for the dollar pass-through: column (3) reports that the a 10% dollar appreciations is associated with an 3.5% increase in import prices. However, our 2SLS estimates show that the impact of the invoicing share on the dollar pass-through is even larger. Column (6) shows that, once the dollar is instrumented by the US housing investment, a 10% increase in the dollar invoicing share is associated with an 11.3% increase in import prices, which is almost 8% more than based on the OLS estimate. More importantly, the coefficient of about one is exactly what the DCP predicts.

As before, we also estimate our specification with the updated data from Boz et al. (2020), which is shown in columns (4) and (7). The estimated interaction term coefficient increases slightly in column (7) but is not statistically different from the coefficient in column (6) at conventional levels.

Robustness Checks In this section, we provide robustness checks for our IV estimation strategy. We begin with robustness results for the import quantity regressions. First, one may be concerned that U.S. domestic conditions may affect countries' imports from other countries through another channel than the dollar exchange rate. If that was the case, one would expect effects to differ across countries depending on their exposure to the United

States. In particular, effects could be larger for countries with closer links to the United States, and one may be worried that effects are driven by a few countries with strong trade links with the United States. To address this issue, we split our sample into importing countries with a below-median and an above-median share of imports from the United States.[8] Table 3 reports the results. Panel A shows OLS and 2SLS estimates for importers with below-median US trade share and Panel B shows estimates for above-median countries. Column (3) in both panels shows that while the dollar appreciation is still associated with a decline in the trading volume for either above-median or below-median importers, the OLS coefficient for the interaction between the dollar exchange rate and the dollar invoicing share is no longer statistically significant at five-percent level in both cases. Turning to the IV estimates reported in columns (4) and (5), we find that the causal link the dollar and trade volume established in Table 1 remains robust and holds in each of the two sub-samples. The interaction term coefficients are highly statistically significant and negative in both panel A and B. Thus, our findings are not driven by importers with large trade exposures to the United States, but also hold for countries with below-median exposure.

In Panel A of Table 4, we provide robustness on the choice of the instrument, adding additional lags of housing permits, and using our two alternative instruments capturing U.S. housing activity. All columns in the table report the second stage results of our 2SLS estimation. The first two columns show that results remain robust when we include up to 3-year lags of building permits as instruments for the dollar exchange rate. The lagged values of building permits control for the high serial correlations of the permits in the first-stage

[8] For each country, we calculate the import share as the total imports from the United States from 1988 to 2015 divided by the total imports from all 53 trade partners in the sample. The median share of imports from the United States in total imports is 1.9%.

regression. Columns (3) - (4) present results when the dollar is instrumented by residential investment (PRFI) scaled by gross investment (GDPI), while columns (5) - (6) report results when the dollar is instrumented by housing starts. We find that both alternative instruments provide similar results compared to our baseline estimations: the interaction terms of the dollar exchange rate with the dollar invoicing share remain statistically significantly in all cases and coefficients are substantially larger than in the OLS baseline.

The 2007-2009 Great Financial Crisis (GFC) ended a great boom and bust cycle for the U.S. housing market. As depicted in Figure 1, US housing permits rose sharply before the GFC and then dropped to a historical low in 2009. Are our results driven by the boom-bust cycle of the U.S. housing market around the GFC? To answer this question, Panel B of Table 4 report 2SLS results that are estimated over several sub-periods of our sample. The first two columns report the estimates when we exclude the GFC years 2007 to 2009 from our sample, columns (3) and (4) present results for the pre-2007 sub-sample, and columns (5) and (6) reports estimates for the post-2009 sub-sample. We find that results for the different sub-samples are quite similar to the baseline results. In all sub-samples, the coefficients for the dollar and the interaction term of the dollar with the invoicing share are negative and statistically significant. Results differ the most for the post-GFC sample, where the interaction term coefficient declines to -0.44 and is only significant at the 5 percent level. However, that sub-sample only covers 6 years, 2010 to 2015, and this result should therefore be interpreted with caution. Importantly, as shown in columns (1) to (4), results are very strong for the sample that excludes the GFC and the pre-2007 sample. Thus, the casual link between the dollar and import quantities we identify is not driven by observations around the GFC.

Tables 5 and 6 provide the same robustness checks discussed above for estimates of the pass-through regression (6). Our 2SLS estimates from Table 5 show that the baseline estimates remain robust when we split our sample into importers with below-median and above-median U.S. import shares. Column (5) in Panels A and B of Table 5 show that the interaction term between the dollar exchange rate and the dollar invoicing share is positive and highly significant in both sub-samples. Compared to the OLS results reported in Column (3), the dollar invoicing share has a larger effect on import prices based on the 2SLS estimates. A 10% increase in the dollar invoicing share is associated with an 13% increase in import prices for importers with below-median trade share with US, compared to 3.7% increase estimated by OLS. For above-median countries, a 10% increase in the dollar invoicing share is associated with an 10.9% increase in import prices, compared to 4.1% from OLS.

Panel A of Table 6 reports results using alternative measures of housing cycles. We continue to find that regressions using the alternative U.S. housing instruments generate similar results to our baseline 2SLS estimation. Panel B reports 2SLS results that are estimated over several sub-periods of our sample. Results for different sub-periods are very similar to the ones obtained from the full sample. In all sub-periods investigated, the coefficient of the dollar and that of the interaction term of the dollar with the invoicing share remain statistically significant and positive. Again, we find weaker results for the post-GFC period. For that period, we estimate an interaction term coefficient of 0.144 that is only significant at the 10 percent level. However, as noted above, the post-GFC sample is quite short and only covers six years from 2010 to 2015.

6 Conclusion

This paper provides evidence on a causal effect of the dollar on international trade, using a new instrument for the dollar that is based on domestic U.S. housing conditions. We find that changes to a country's exchange rate vis-a-vis the dollar are fully passed through to import prices and that trade quantities decline notably with dollar appreciation. Both findings lend strong support to the DCP proposed in Gopinath et al. (2020). The substantially larger coefficient estimates obtained through our 2SLS estimation indicate that the endogeneity of the dollar is a problem in the baseline OLS results. Reassuringly, the new larger coefficients are much closer to the theoretical predictions in Gopinath et al. (2020), underscoring the plausibility of our approach.

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Table 1. Trade Elasticity with respect to Exchange Rate

	OLS				IV		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
S_j source	Gopinath et al. (2020)			Boz et al. (2020)	Gopinath et al. (2020)		Boz et al. (2020)
$\Delta e_{ij,t}$	-0.119***	-0.031*	-0.077*	-0.141**	0.394***	-0.320***	-0.221***
	(0.0139)	(0.0160)	(0.0403)	(0.0581)	(0.0778)	(0.0632)	(0.0847)
$\Delta e_{ij,t} \times S_j$			0.118*	0.245**		1.296***	2.062***
			(0.0684)	(0.1001)		(0.2420)	(0.414)
$\Delta e_{\$j,t}$		-0.186***	-0.140**	-0.064	-1.052***		
		(0.0250)	(0.0600)	(0.0807)	(0.161)		
$\Delta e_{\$j,t} \times S_j$			-0.090	-0.330**		-1.522***	-2.007***
			(0.0871)	(0.1323)		(0.256)	(0.816)
First Stage F-Stats					990.55	2253.30	1193.22
Observations	52,272	52,272	38,582	28,474	52,272	38,582	28,474
Number of dyad	2,807	2,807	2,014	2,258	2,807	2,014	2,258

All regressions include two lags of the independent variables, lags 0-2 of importer real GDP growth, and time fixed-effects. The standard errors in column (1) - (4) are clustered by dyads and associated standard errors are reported in parenthesis. Column (5) reports the 2SLS estimates using one-year lag of U.S. building permits growth as the instrument for the changes in US dollar. Column (6) and (7) reports the 2SLS estimates using one-year lag of U.S. building permits growth $\times$ Dollar-invoicing Shares S_j as the instrument for the changes in US dollar $\times$ Dollar-invoicing Shares. The dollar invoicing share in column (3) and (6) are from from Gopinath et al. (2020), and those in column (4) and (7) are from Boz et al. (2020). The standard errors are reported in parenthesis. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$

Table 2. Exchange Rate Pass-through into Prices

	OLS				IV		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
S_j source	Gopinath et al. (2020)			Boz et al. (2020)	Gopinath et al. (2020)		Boz et al. (2020)
$\Delta e_{ij,t}$	0.757**	0.164***	0.209***	0.276***	0.347***	0.422***	0.506***
	(0.0132)	(0.0126)	(0.0169)	(0.0260)	(0.0216)	(0.0227)	(0.0339)
$\Delta e_{ij,t} \times S_j$			-0.0841***	-0.222***		-0.378***	-0.718***
			(0.0240)	(0.0470)		(0.0617)	(0.102)
$\Delta e_{\$j,t}$		0.781***	0.565***	0.620***	0.588***		
		(0.0143)	(0.0283)	(0.0402)	(0.0319)		
$\Delta e_{\$j,t} \times S_j$			0.348***	0.281***		1.133***	1.304***
			(0.0326)	(0.0595)		(0.0689)	(0.127)
First-stage F-Stats					4870.45	2435.06	617.08
Observations	46,820	46,820	34,513	25,597	46,820	34,513	25,597
Number of dyad	2,647	2,647	1,900	2,112	2,647	1,900	2,112

All regressions include two lags of the independent variables, lags 0-2 of exporter changes in producer price index, and time fixed-effects. The standard errors in column (1) - (4) are clustered by dyads and associated standard errors are reported in parenthesis. Column (5) reports the 2SLS estimates using one-year lag of U.S. building permits growth as the instrument for the changes in US dollar. Column (6) and (7) reports the 2SLS estimates using one-year lag of U.S. building permits growth $\times$ Dollar-invoicing Shares S_j as the instrument for the changes in US dollar $\times$ Dollar-invoicing Shares. The dollar invoicing share in column (3) and (6) are from from Gopinath et al. (2020), and those in column (4) and (7) are from Boz et al. (2020). The standard errors are reported in parenthesis. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$

Table 3. Robustness: Trade Elasticity with respect to Exchange Rate

Panel A: Countries with **below-median** trade share with US					
	OLS			IV	
	(1)	(2)	(3)	(4)	(5)
$\Delta e_{ij,t}$	-0.0368*	-0.00463	-0.0161	0.273***	0.0197
	(0.0206)	(0.0219)	(0.0531)	(0.0511)	(0.0551)
$\Delta e_{ij,t} \times S_j$			0.0346		0.241**
			(0.0934)		(0.114)
$\Delta e_{\$j,t}$		-0.140***	-0.162*	-1.284***	
		(0.0456)	(0.0912)	(0.180)	
$\Delta e_{\$j,t} \times S_j$			-0.0926		-1.736***
			(0.183)		(0.275)
First-stage F-Stats				936.88	962.47
Observations	23,375	23,375	20,972	23,375	20,972
Number of dyad	1,208	1,208	1,052	1,208	1,052
Panel B: Countries with **above-median** trade share with US					
	(1)	(2)	(3)	(4)	(5)
$\Delta e_{ij,t}$	-0.159***	-0.0548**	-0.186***	0.249*	-0.283***
	(0.0178)	(0.0230)	(0.0607)	(0.132)	(0.0762)
$\Delta e_{ij,t} \times S_j$			0.258***		0.669**
			(0.0974)		(0.263)
$\Delta e_{\$j,t}$		-0.178***	-0.123	-0.673***	
		(0.0311)	(0.0899)	(0.226)	
$\Delta e_{\$j,t} \times S_j$			-0.121		-0.619***
			(0.123)		(0.236)
First-stage F-Stats				2133.53	1518.79
Observations	28,897	28,897	17,610	28,897	17,610
Number of dyad	1,599	1,599	962	1,599	962

All regressions include two lags of the independent variables, lags 0-2 of importer real GDP growth, and time fixed-effects. The standard errors in column (1) - (3) are clustered by dyads and associated t-stats are reported in parenthesis. Column (4) reports the 2SLS estimates using one-year lag of U.S. building permits growth as the instrument for the changes in US dollar. Column (5) reports the 2SLS estimates using one-year lag of U.S. building permits growth × Dollar-invoicing Shares S_j from Gopinath et al. (2020) as the instrument for the changes in US dollar × Dollar-invoicing Shares . The standard errors are reported in parenthesis. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$

Table 4. Robustness: Trade Elasticity with respect to Exchange Rate

Panel A: Alternative Instruments						
Instruments:	1-3 lags of Permits		US PRFI/GDPI		Housing Starts	
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta e_{ij,t}$	0.0739*	-0.233***	0.208***	-0.287***	0.476***	-0.354***
	(0.0383)	(0.0517)	(0.0524)	(0.0565)	(0.0883)	(0.0685)
$\Delta e_{ij,t} \times S_j$		0.831***		1.120***		1.477***
		(0.172)		(0.199)		(0.271)
$\Delta e_{\$j,t}$	-0.383***		-0.665***		-1.225***	
	(0.0728)		(0.104)		(0.183)	
$\Delta e_{\$j,t} \times S_j$		-0.991***		-1.321***		-1.729***
		(0.170)		(0.203)		(0.288)
First-stage F-Stats	2758.33	2779.72	2943.02	2261.94	2906.40	2241.23
Observations	52,272	38,582	52,272	38,582	52,272	38,582
Number of dyad	2,807	2,014	2,807	2,014	2,807	2,014
Panel B: Sub-periods						
Sub-periods	Exclude GFC		Pre–GFC		Post-GFC	
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta e_{ij,t}$	0.170***	-0.337***	0.140***	-0.316***	0.245***	0.467**
	(0.0538)	(0.0742)	(0.0491)	(0.0509)	(0.0683)	(0.182)
$\Delta e_{ij,t} \times S_j$		1.193***		0.961***		-0.535*
		(0.291)		(0.157)		(0.273)
$\Delta e_{\$j,t}$	-0.640***		-0.587***		-0.294***	
	(0.109)		(0.102)		(0.0695)	
$\Delta e_{\$j,t} \times S_j$		-1.351***		-1.038***		-0.442**
		(0.300)		(0.158)		(0.181)
First-stage F-Stats	1448.84	1073.43	776.05	639.42	237.08	161.24
Observations	35,691	26,644	36,050	26,841	13,464	9,755
Number of dyad	2,805	2,014	2,779	1,992	2,791144	2,010

All regressions include two lags of the independent variables, lags 0-2 of importer real GDP growth, and time fixed-effects. The standard errors in column (1) - (3) are clustered by dyads and associated standard errors are reported in parenthesis. Column (4) reports the 2SLS estimates using one-year lag of U.S. building permits growth as the instrument for the changes in US dollar. Column (5) reports the 2SLS estimates using one-year lag of U.S. building permits growth $\times$ Dollar-invoicing Shares S_j from Gopinath et al. (2020) as the instrument for the changes in US dollar $\times$ Dollar-invoicing Shares. PRFI/GDPI is the share of US gross domestic private investment (GDPI) attributable to the private residential fixed investment (PRFI). GFC stands for 2007-2009 Great Financial Crisis. The standard errors are reported in parenthesis. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$

Table 5. Robustness: Exchange Rate Pass-through into Price

Panel A: Countries with **below-median** trade share with US					
	OLS			IV	
	(1)	(2)	(3)	(4)	(5)
$\Delta e_{ij,t}$	0.518***	0.160***	0.182***	0.287***	0.304***
	(0.0191)	(0.0159)	(0.0205)	(0.0223)	(0.0241)
$\Delta e_{ij,t} \times S_j$			-0.0662**		-0.0895**
			(0.0301)		(0.0390)
$\Delta e_{\$j,t}$		0.724***	0.588***	0.577***	
		(0.0252)	(0.0406)	(0.0498)	
$\Delta e_{\$j,t} \times S_j$			0.376***		1.304***
			(0.0635)		(0.127)
First-stage F-Stats				1715.29	2369.57
Observations	20,890	20,890	18,714	20,890	18,714
Number of dyad	1,144	1,144	996	1,144	996
Panel B: Countries with **above-median** trade share with US					
	(1)	(2)	(3)	(4)	(5)
$\Delta e_{ij,t}$	0.826***	0.171***	0.240***	0.366***	0.478***
	(0.0124)	(0.0192)	(0.0289)	(0.0339)	(0.0372)
$\Delta e_{ij,t} \times S_j$			-0.0893**		-0.440***
			(0.0378)		(0.0932)
$\Delta e_{\$j,t}$		0.785***	0.489***	0.588***	
		(0.0203)	(0.0426)	(0.0445)	
$\Delta e_{\$j,t} \times S_j$			0.409***		1.093***
			(0.0459)		(0.0914)
First-stage F-Stats				1545.55	3359.83
Observations	25,930	25,930	15,799	25,930	15,799
Number of dyad	1,503	1,503	904	1,503	904

All regressions include two lags of the independent variables, lags 0-2 of exporter log changes in PPI, and time fixed-effects. The standard errors in column (1) - (3) are clustered by dyads and associated standard errors are reported in parenthesis. Column (4) reports the 2SLS estimates using one-year lag of U.S. building permits growth as the instrument for the changes in US dollar. Column (5) reports the 2SLS estimates using one-year lag of U.S. building permits growth × Dollar-invoicing Shares S_j from Gopinath et al. (2020) as the instrument for the changes in US dollar × Dollar-invoicing Shares . The standard errors are reported in parenthesis. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$

Table 6. Robustness: Exchange Rate Pass-through into Price

Panel A: Alternative Instruments						
Instruments:	1-3 lags of permits		US PRFI/GDPI		Housing Starts	
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta e_{ij,t}$	0.460***	0.432***	0.487***	0.436***	0.348***	0.422***
	(0.0149)	(0.0183)	(0.0164)	(0.0177)	(0.0226)	(0.0227)
$\Delta e_{ij,t} \times S_j$		-0.122***		-0.0243		-0.377***
		(0.0405)		(0.0429)		
$\Delta e_{\$j,t}$	0.417***		0.376***		0.586***	
	(0.0194)		(0.0216)		(0.0335)	
$\Delta e_{\$j,t} \times S_j$		0.788***		0.656***		1.132***
		(0.0454)		(0.0520)		(0.0735)
First-stage F-Stats	886.86	2911.48	3138.89	2831.91	1396.95	2645.69
Observations	46,820	34,513	46,820	34,513	46,820	34,513
Number of dyad	2,647	1,900	2,647	1,900	2,647	1,900
Panel B: Sub-periods						
Sub-periods	Exclude GFC		Pre–GFC		Post-GFC	
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta e_{ij,t}$	0.485***	0.456***	0.420***	0.415***	0.461***	0.353***
	(0.0232)	(0.0216)	(0.0456)	(0.0262)	(0.0240)	(0.0452)
$\Delta e_{ij,t} \times S_j$		0.0525		-0.190***		0.160*
		(0.0576)		(0.0643)		(0.0895)
$\Delta e_{\$j,t}$	0.405***		0.512***		0.127***	
	(0.0288)		(0.0575)		(0.0369)	
$\Delta e_{\$j,t} \times S_j$		0.541***		0.913***		0.144*
		(0.0774)		(0.0883)		(0.0867)
First-stage F-Stats	889.01	2431.71	1213.38	3017.44	1115.21	804.85
Observations	31,912	23,787	29,803	22,194	12,061	8,750
Number of dyad	2,645	1,900	2,522	1,808	2,578	1,858

All regressions include two lags of the independent variables, lags 0-2 of exporter log changes in PPI, and time fixed-effects. The standard errors in column (1) - (3) are clustered by dyads and associated standard errors are reported in parenthesis. Column (4) reports the 2SLS estimates using one-year lag of U.S. building permits growth as the instrument for the changes in US dollar. Column (5) reports the 2SLS estimates using one-year lag of U.S. building permits growth × Dollar-invoicing Shares S_j from Gopinath et al. (2020) as the instrument for the changes in US dollar × Dollar-invoicing Shares . The standard errors are reported in parenthesis. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$