Global Trade and the Dollar*

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Abstract: We document that the U.S. dollar exchange rate drives global trade prices and volumes. Using a newly constructed data set of bilateral price and volume indices for more than 2,500 country pairs, we establish the following facts: 1) Bilateral non-commodities terms of trade are essentially uncorrelated with bilateral exchange rates. 2) The dollar exchange rate quantitatively dominates the bilateral exchange rate in price pass-through and trade elasticity regressions. 3) A 1% U.S. dollar appreciation against all other currencies in the world predicts a 0.6% decline within a year in the volume of total trade between countries in the rest of the world, controlling for the global business cycle. 4) Using a Bayesian semiparametric hierarchical panel data model, we estimate that the importing country's share of imports invoiced in dollars explains 15% of the variance of dollar pass-through across country pairs. Our findings strongly support the dominant currency paradigm as opposed to the traditional Mundell-Fleming pricing paradigms. We then employ a three country model with dollar pricing to demonstrate the asymmetries between the transmission of monetary policy shocks that arise in the U.S. and the rest of the world.

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1 Introduction

Exchange rate fluctuations impact a country's trade competitiveness, inflation, and output and therefore have important consequences for its welfare and economic policy. Friedman (1953) famously championed the virtues of a flexible exchange rate policy because, he argued, it generates just the "right" changes in a country's import and export prices to keep the economy at full employment even when prices are sticky in the producer's currency. This insight is also a central prediction of the canonical Mundell-Fleming paradigm in international macroeconomics (Mundell, 1963; Fleming, 1962; Obstfeld and Rogoff, 1995; Clarida et al., 2001; Galí and Monacelli, 2005). Specifically, this paradigm predicts that, because prices are sticky in the producer's currency, a nominal exchange rate depreciation is associated with a *depreciation* of a country's terms of trade, that is, the ratio of the price of its imports to that of its exports increases when the nominal exchange rate depreciates. A second influential paradigm by Betts and Devereux (2000) and Devereux and Engel (2003) assumes instead that prices are sticky in local (destination country's) currency and therefore has at its core the opposite prediction, namely that a nominal exchange rate depreciation is associated with an *appreciation* of a country's terms of trade, implying different normative recommendations. A common feature of both paradigms is that a country's exchange rate is only as important as its share in world trade, with no exchange rate having a central role.

There is growing evidence, though, that the vast majority of invoicing is neither in the local currency or in the producer's currency but instead in a "dominant currency", which is most often the U.S. dollar (Goldberg and Tille, 2008; Gopinath, 2015). Consistent with this evidence, Casas et al. (2016) develop a "dominant currency paradigm" (DCP) that assumes that prices are mostly sticky in the dollar.[1] This paradigm predicts that the terms of trade are only *weakly sensitive* to the exchange rate and the value of a country's currency relative to the dollar is a primary driver of a country's import prices and quantities regardless of where the good originates from. That is, the dollar has a central role in world trade. Importantly, these differences in predictions across pricing paradigms arise not only when the currency of invoicing is exogenously imposed, but also when it is the outcome of an endogenous decision by firms.

In this paper we investigate the empirical validity of the local, producer, and dominant currency pricing paradigms in international macroeconomics. To do so, we construct harmonized annual *bilateral* import and export unit value and volume indices for 55 countries, yielding more

[1] Other references on dollar pricing include Corsetti and Pesenti (2005), Goldberg and Tille (2009), Devereux et al. (2007), and Canzoneri et al. (2013). Devereux and Engel (2007) analyze the case of producer currency pricing for at-the-dock prices and local currency pricing for consumer prices.

than 2,500 dyads, i.e., trading pairs. The indices are constructed from highly disaggregated UN Comtrade customs data starting as early as 1989, depending on the country, and covering through 2015. The countries in our sample comprise 91% of the world's total goods exports and imports in 2015. Importantly, we exclude commodities from these indices because the paradigms are relevant only for goods with sticky prices.

We document four facts that pose a serious challenge to the mainstream paradigms but support DCP: 1) The terms of trade neither depreciate nor appreciate alongside a depreciation of the nominal exchange rate. They are best described as being insensitive to it. 2) The U.S. dollar has a disproportionate impact on world trade: The dollar exchange rate quantitatively dominates the bilateral exchange rate in price pass-through and trade elasticity regressions, even when the U.S. is on neither side of the trade transaction. 3) The strength of the U.S. dollar is a key predictor of rest-of-world aggregate trade volume and consumer/producer price inflation. Specifically, a 1% U.S. dollar appreciation against all other currencies in the world predicts a 0.6% decline within a year in the volume of total trade between countries in the rest of the world, holding constant proxies for the global business cycle. 4) The importer's share of imports invoiced in dollars explains 15% of the variance of dollar pass-through across country pairs. Our results derive from fixed effects panel regressions as well as a novel Bayesian semiparametric hierarchical panel data model. Lastly, given the empirical support for DCP, we employ a three country model to contast the transmission of monetary policy shocks arising in the U.S. versus in the rest of the world.

We now elaborate further on each of our findings. For our first finding on the terms of trade, we regress changes in the bilateral non-commodities terms of trade on changes in the bilateral exchange rate. This yields a contemporaneous coefficient on the exchange rate of 0.037, with the 95% confidence interval $[0.02, 0.05]$. For comparison, the coefficient should be close to 1 under producer currency pricing and to -1 under local currency pricing.[2]

For our second finding, we estimate exchange rate pass-through and trade elasticity regressions at the country pair level. We first follow standard practice and estimate the pass-through of *bilateral* exchange rates into import prices and volumes. This practice follows naturally from the classic Mundell-Fleming paradigm, according to which the price an importing country faces (when expressed in the importing country's currency) fluctuates closely with the bilateral exchange rate. Accordingly, studies of exchange rate pass-through focus on trade-weighted or bilateral exchange rate changes (Goldberg and Knetter, 1997; Burstein and Gopinath, 2014). In standard price pass-through regressions, we document that when country j's currency depreci-

[2]As we explain later, this finding cannot be explained by a model with flexible prices and strategic complementarities in pricing as in Atkeson and Burstein (2008) and Itskhoki and Mukhin (2017).

ates relative to country i by 10%, import prices in country j for goods imported from country i rise by 8%, suggestive of close to complete pass-through at the one year horizon. Next, we introduce the U.S. dollar exchange rate as an additional explanatory variable in price pass-through and trade elasticity regressions. Adding the dollar to the regression knocks the coefficient on the bilateral exchange rate from 0.76 down to 0.16. The coefficient on the dollar exchange rate of 0.78 dominates that of the bilateral exchange rate. This result cannot be attributed to the correlation of the dollar with the global business cycle, as we control for time fixed effects. There is also no evidence of collinearity across exchange rates because all the coefficients are precisely estimated. Moreover we show that the magnitude of the dollar pass-through is indeed systematically related to the dollar invoicing shares of countries. Specifically, increasing the dollar invoicing share by 10 percentage points causes the contemporaneous dollar pass-through to increase by 3.5 percentage points. Similar to the price regressions, adding the dollar exchange rate to a bilateral volume forecasting regression knocks down the coefficient on the bilateral exchange rate by a substantial amount. The contemporaneous volume elasticity for the dollar exchange rate is -0.19, while the elasticity for the bilateral exchange rate is an order of magnitude smaller at -0.03.

These pass-through estimates point to a potential misspecification in standard pass-through regressions that ignore the role of the dollar. We also show that the dollar's role as an invoicing currency is indeed special, as it handily beats the explanatory power of the euro in price and volume regressions. The data is also consistent with an additional key prediction of the dominant currency paradigm: U.S. import volumes are significantly less sensitive to the bilateral exchange rate, as compared to other countries' imports.

Third, we demonstrate empirically that the strength of the U.S. dollar is a key predictor of rest-of-world trade volume and inflation, again controlling for measures of the global business cycle. We find that a 1% appreciation of the U.S. dollar relative to all other currencies is associated with a 0.6% contraction in rest-of-world (i.e., excluding the U.S.) aggregate import volume within the year. Furthermore, countries with larger dollar import invoicing shares experience higher pass-through of the dollar exchange rate into consumer and producer price inflation.

Fourth, we exploit our rich panel data set to show that the cross-dyad heterogeneity in pass-through coefficients is well explained by the propensity to invoice imports in dollars. We use the importer's country-level dollar invoicing share from Gopinath (2015) as a proxy for the invoicing share of bilateral imports.[3] Using a flexible hierarchical Bayesian framework to directly model pass-through heterogeneity, we estimate that the importer's dollar invoicing share explains 15%

[3] Casas et al. (2016) use customs data to calculate export invoicing shares for Colombia at the bilateral level and find small heterogeneity, implying that a country's average should serve as a good proxy for some countries.

of the overall cross-dyad variance in dollar exchange rate pass-through into prices. We also find that the importer's dollar invoicing share affects the exchange rate elasticity of trade volumes. These findings confirm the quantitative importance of the global currency of invoicing, a key concept in the dominant currency paradigm.

Our Bayesian estimation procedure allows the data to speak flexibly about the extent and determinants of the cross-sectional heterogeneity of pass-through. We employ a random coefficients panel data model, where the distribution of the random coefficient on the dollar exchange rate is allowed to depend nonparametrically on the dollar invoicing share as well as other unobserved determinants of pass-through heterogeneity. Unlike standard panel regressions that are informative about the *average* pass-through and the statistical significance of the determinants of pass-through heterogeneity, the Bayesian approach allows us to quantify the *overall* cross-sectional heterogeneity of exchange rate pass-through/elasticities *and* the relation of this heterogeneity to dollar invoicing. The hierarchical aspect of the Bayesian inference procedure can be thought of as striking an optimal bias-variance balance between two extreme approaches: (1) constrained but precisely estimated panel regressions with cross-sectionally constant parameters and (2) unconstrained but noisily estimated dyad-by-dyad time series regressions.

Given our findings, we conclude that DCP is a more empirically relevant starting point than traditional modeling approaches and provide a DCP framework to analyze the international transmission of shocks. Since our findings suggest a special role for U.S. monetary policy we simulate a DCP model with three large countries and demonstrate that, when monetary policy is set using a Taylor rule: (i) monetary policy (MP) shocks in the U.S. have strong spillovers to MP in other countries, while the reverse is not true; (ii) A tighter MP in the U.S. and the accompanying appreciation of the dollar has a significant negative impact on rest-of-world and global trade, while this is not true for monetary tightenings originating in other countries.

Literature review. Our work on the terms of trade is related to Obstfeld and Rogoff (2000), who conduct one of the earliest tests of the Mundell-Fleming paradigm against the Betts-Devereux-Engel paradigm. Obstfeld and Rogoff (2000) examine the *correlation* between *country*-level terms of trade and the trade-weighted exchange rate for 21 countries, using quarterly data for 1982-1998. They report an average correlation of 0.26, which they interpret as a rejection of local currency pricing. Even though the correlation is well less than 1, which would lend weak support for producer currency pricing, they conjecture that the low correlation could be because of the construction of the trade-weighted exchange rates and/or because their terms of trade measures include commodity prices.[4] With the help of our globally representative data set, we improve upon Obstfeld and Rogoff (2000) in several dimensions. Specifically, we examine the *bilateral* terms of trade, excluding commodity prices. We estimate pass-through coefficients as opposed to correlations, as a high correlation alone is not sufficient to support producer currency pricing, which predicts a high pass-through. Moreover, in addition to studying the relationship between the terms of trade and the exchange rate, we test a battery of predictions by the different pricing paradigms.

Our exchange rate pass-through analysis appears to be among the first to exploit a globally representative data set on bilateral trade volumes in addition to values. This allows us to distinguish the effects of exchange rates on volumes and prices (more precisely, unit values) at the level of country pairs. We use the cross-sectional richness of our data set to investigate the determinants of differential pass-through, especially as it relates to currency of invoicing. To our knowledge, the only other work that utilizes a similarly rich data set is Bussière et al. (2016), who analyze trade prices and quantities at the product level. The goal of that paper is to quantify the elasticity of prices and quantities to the bilateral exchange rate and check if Marshall-Lerner conditions hold. In contrast, our goal is to empirically evaluate the predictions of the various pricing paradigms and in the process highlight the dollar's central role in global trade.

The remaining literature on exchange rate pass-through falls into two main camps. First, many papers use unilateral (i.e., country-level) time series, which limits the ability to analyze cross-sectional heterogeneity and necessitates the use of trade-weighted rather than truly bilateral exchange rates (e.g., Leigh et al., 2015). Second, a recent literature estimates pass-through of bilateral exchange rates into product-level prices, as opposed to unit values, but these micro data sets are available for only a few countries (see the review by Burstein and Gopinath, 2014).

Our paper confirms that the findings of Casas et al. (2016) are relevant for the majority of world trade, and we establish additional channels of U.S. dollar dominance. Casas et al. (2016) model DCP for a small open economy and test its implications for the dollar's role in pass-through regressions and for the terms of trade using product level data for Colombia. We depart from this paper by providing evidence for DCP using newly constructed import and export indices for 55 countries and over 2,500 country pairs. Moreover, we relate the heterogeneity in pass-through estimates to the dollar invoicing shares of countries and demonstrate that the strength of the dollar is a key predictor of rest-of-world trade and consumer/producer price inflation. In addition,

[4] To quote, they say (p. 124): "First, the IMF nominal effective exchange rate measure (...) covers industrial trading partners only, with weights based on manufacturing trade, while the terms of trade index covers all trading partners and all goods. Second, some of the goods entering the terms of trade index, especially on the import side, are flexible-price commodities."

we employ a DCP model with three large countries to examine the asymmetries in transmission of monetary policy shocks arising in the U.S. versus those arising in the rest of the world.

From a methodological perspective, our paper's contribution is to leverage a semiparametric Bayesian panel data model with cross-sectionally heterogeneous slope coefficients. Our specification of the cross-sectional distribution of slope coefficients relies on the Mixture of Gaussian Linear Regression (MGLR) nonparametric conditional density prior in Pati et al. (2013), who derive high-level posterior concentration results. The MGLR prior extends the much-used Dirichlet Process Mixture prior to *conditional* density estimation. Liu (2017) uses a similar MGLR prior specification for semiparametric Bayesian panel data modeling, but she focuses on forecasting rather than characterizing cross-sectional heterogeneity. Although our linear-in-parameters specification is more restrictive than the frequentist non-parametric approaches of Evdokimov (2010) and Chernozhukov et al. (2013), our Bayesian framework facilitates visualization of the conditional pass-through distribution, uncertainty assessment, and model selection.

Outline. Section 2 discusses a simple conceptual framework that guides our empirical analysis and interpretations. Section 3 describes our data set of bilateral trade unit values and quantities, exchange rates, and dollar invoicing shares. Section 4 presents panel regression evidence on the average pass-through from bilateral, dollar, and euro exchange rates into prices, quantities, and terms of trade; moreover, it quantifies the role of the U.S. dollar as a predictor of rest-of-world trade volume and inflation. In Section 5, we employ the Bayesian model to characterize the cross-sectional heterogeneity of dollar pass-through and its relation to invoicing shares. Section 6 analyzes a structural DCP model to contrast the impact of MP shocks originating from the US versus the rest of the world. Section 7 concludes. Appendix A details the data set and Bayesian approach. Appendix B contains supplementary material on data, empirics, numerical procedures, and the theoretical model.

2 Conceptual framework

In this section we provide a simple conceptual framework along the lines of Proposition 2 in Casas et al. (2016) to motivate the empirical analysis that follows. Define p_{ij} to be the log price of goods exported from country i to country j measured in currency j, e_{ij} to be the log bilateral exchange rate between country i and country j expressed as the price of currency i in terms of currency j, and $e_{\$j}$ to be the log price of a U.S. dollar in currency j. Suppose a fraction θ_i of these exports are invoiced in the producer's (country i) currency, a fraction θ_j in the local (destination country j) currency and a fraction θ_u in the dominant currency (dollar) with $\sum_{k\in\{i,j,u\}} \theta_k = 1$. Import price inflation for country j for goods originating from country i is then

$$\Delta p_{ij,t} = \theta_j \Delta p^j_{ij,t} + \theta_i \left[\Delta p^i_{ij,t} + \Delta e_{ij,t}\right] + \theta_u \left[\Delta p^u_{ij,t} + \Delta e_{\$j,t}\right],$$

where $p^k_{ij,t}$ stands for the price of goods imported by country j from i that are invoiced in currency k. Calvo pricing implies $\Delta p^k_{ij,t} = (1-\delta_p)\left(\bar{p}^k_{ij,t} - p^k_{ij,t-1}\right)$ where $\bar{p}^k_{ij,t}$ is the reset-price for (ij) in currency k. Through substitution we can express import price inflation as

$$\Delta p_{ij,t} = \theta_i \Delta e_{ij,t} + \theta_u \Delta e_{\$j,t} + (1-\delta_p)\sum_k \theta_k \Delta \bar{p}^k_{ij,t},$$

where $\Delta \bar{p}^k_{ij,t} \equiv \bar{p}^k_{ij,t} - p^k_{ij,t-1}$. In the very short run when $\delta_p \to 1$[5], we have the following benchmarks for the changes in import prices and the terms of trade (TOT).

- In the case of producer currency pricing (PCP), $\theta_i = 1$ and $\theta_j = \theta_u = 0$,

$$\begin{aligned} \Delta p_{ij,t} &= 1\cdot\Delta e_{ij,t} + 0\cdot\Delta e_{\$j,t}, \qquad \Delta p_{ji,t} = -1\cdot\Delta e_{ij,t} + 0\cdot\Delta e_{\$j,t}, \\ tot_{ij,t} &= \Delta p_{ij,t} - (\Delta p_{ji,t} + \Delta e_{ij,t}) = 1\cdot\Delta e_{ij,t}. \end{aligned}$$

- In the case of local currency pricing (LCP), $\theta_j = 1$ and $\theta_i = \theta_u = 0$,

$$\begin{aligned} \Delta p_{ij,t} &= 0\cdot\Delta e_{ij,t} + 0\cdot\Delta e_{\$j,t} \qquad \Delta p_{ji,t} = 0\cdot\Delta e_{ij,t} + 0\cdot\Delta e_{\$j,t}, \\ tot_{ij,t} &= \Delta p_{ij,t} - (\Delta p_{ji,t} + \Delta e_{ij,t}) = -1\cdot\Delta e_{ij,t}. \end{aligned}$$

- In the case of dominant currency pricing (DCP), $\theta_u = 1$ and $\theta_i = \theta_j = 0$,

$$\begin{aligned} \Delta p_{ij,t} &= 0\cdot\Delta e_{ij,t} + 1\cdot\Delta e_{\$j,t}, \qquad \Delta p_{ji,t} = 0\cdot\Delta e_{ij,t} + 1\cdot\Delta e_{\$i,t}, \\ tot_{ij,t} &= \Delta p_{ij,t} - (\Delta p_{ji,t} + \Delta e_{ij,t}) = 0\cdot\Delta e_{ij,t}. \end{aligned}$$

[5]A primitive assumption of Keynesian environments is that prices are sticky in the currency of invoicing. This assumption is supported by direct measures of price stickiness for U.S. prices-at-the dock in Gopinath and Rigobon (2008) and for Irish prices-at-the dock in Fitzgerald and Haller (2013). Cravino (2017) provides indirect evidence of price stickiness using differential sensitivity of Chilean export prices invoiced in different currencies to exchange rate shocks, similar to the evidence in Gopinath et al. (2010).

These predictions, when prices are yet to change, do not depend on what drives the exchange rate variation, that is if it arises from monetary policy shocks or financial shocks.[6]

The data we employ is at an annual frequency so it is natural to ask how these predictions hold at that frequency. In this regard we make the following points: As the horizon increases, the frequency of price adjustment increases and the pass-through predictions depend also on the response of reset prices $\bar{p}^k_{ij,t}$ to exchange rates. Casas et al. (2016) demonstrate using simulated data from a model with exogenous invoicing, strategic complementarities in pricing and imported input use that the divergent predictions across the different paradigms hold at an annual frequency. Secondly, with endogenous currency choice, Gopinath et al. (2010) demonstrate that firms choose to price in currencies in which their reset prices are most stable, i.e., desired medium-run pass-through into the price (expressed in the invoicing currency) is low, as this minimizes the distance between their sticky price and their desired price. In other words, $\Delta\bar{p}^k_{ij,t}$ is relatively insensitive to exchange rate changes and this is an unconditional relation in the sense of not depending on the source of exchange rate movement, in standard models. Consequently, even at longer horizons, we expect the divergent predictions across PCP, LCP, and DCP to hold. That is, the testable differences across the different pricing paradigms are only reinforced when currency of invoicing is an endogenous decision.

In summary, the implications for the terms of trade and the predicted coefficients from a regression of (log) changes in import prices on (log) changes in bilateral and dollar exchange rates differ qualitatively across the PCP, LCP, and DCP benchmarks. Therefore, these regressions provide a useful lens through which to investigate the validity of the three pricing paradigms. While countries are of course not necessarily at corners with regard to their pricing regimes, we should still expect those that rely more heavily on dollar pricing to display greater sensitivity to the dollar exchange rate, even when controlling for the bilateral exchange rate. This is another testable prediction that we investigate. Section 6 lays out a fully dynamic three country model with DCP that confirms several of the predictions highlighted here.

3 Data

The core of our data set consists of panel data on bilateral trade values and volumes from Comtrade. To this global data set we append macroeconomic country aggregates from the World

[6]Note that if the source of the shock generates a strong co-movement across exchange rates, that is there is collinearity, then it would show up in the regressions as large standard errors around the point estimates. As we show, this is not an issue.

Bank's World Development Indicators, financial variables from the Federal Reserve Bank of St. Louis's FRED database, and currency invoicing shares from Gopinath (2015).

Comtrade. UN Comtrade provides detailed annual customs data for a large set of countries at HS 6-digit product level with information about the destination country, USD value, quantity, and weight of imports and exports. This dataset makes it possible to compute volume changes over time for each product, and use the value data to infer unit values. Once unit values are calculated, we compute chained Fisher price indices to aggregate up from the product level to the bilateral country level.[7] We focus entirely on data for non-commodity goods, except noted otherwise. Given the inherent difficulty in drawing a line between commodities and non-commodities, we define commodities fairly broadly as HS chapters 1–27 and 72–83, which comprise animal, vegetable, food, mineral, and metal products.

Coverage of Comtrade at annual frequency over time and across countries is good. The longest time span of the data is 1989–2015, although the coverage varies by dyad. Appendix A.1 lists the coverage by country. In 2015, the 55 countries in our sample were responsible for 91.2% and 91.5% of the value of world goods imports and exports, respectively, as recorded in Comtrade. We exclusively use Comtrade data reported by the importing country, as importer-reported data is regarded as being more reliable since imports generate tariff revenues (Feenstra et al., 2005; World Bank, 2010).

The biggest challenge for constructing price and volume indices using customs data is the so-called unit value bias, as argued by Silver (2007). Unit values, calculated simply by dividing observed values by quantities, are not actual prices. Even at the narrowly defined product categories at 6-digit product level, there is likely to be a wide range of products whose prices may not be moving proportionately. The implication is that if there are shifts in quantities traded within the narrowly defined product categories, unit values would be influenced even when there is no price movement. This creates a bias that the employed methodology takes a stab at correcting for by eliminating products whose unit values have a variance higher than a threshold and are more likely to be biased.

Another challenge that arises from using Comtrade data is related to the use of different HS vintages over time. HS classification is updated about every five years to ensure that the available codings accurately reflect the variety of products being traded. This involves introducing codes for new products, eliminating the old ones, and often regrouping existing products. While con-

[7] The Fisher price index satisfies a number of tests laid out in index number theory and is flexible enough to provide a good proxy for a large set of functional forms (Gaulier et al., 2008; IMF, 2009).

cordances are readily available to facilitate the matching of HS codes across different HS vintages, this process inevitably leads to a loss of information, especially in the case of data on quantities, because the mapping of products across vintages is rarely one-to-one. To get around this problem, for the years in which there is a transition to a new HS vintage, we compute the indices twice, once under the old vintage (using concordances) and once under the new one. This way, only these transition years would be effected by the loss of information due to matching across vintages. After that year, we switch to working with the new vintage. This method not only minimizes the loss of information but also allows us to include new products in the construction of the indices. Boz and Cerutti (2017) provide further details of this method, including the strategy for dealing with outliers and missing values, and a comparison with a similar dataset constructed by Gaulier et al. (2008).

In the final stage, we compare our unit value indices to those provided by the Bureau of Labor Statistics (BLS) for the U.S., the only country, to our knowledge, that collects import price indices based on price surveys by origin. As shown in Appendix B.1, this comparison for the U.S. suggests that working with unit values is acceptable, as the growth rates of the two series are broadly aligned for most trading partners. Further, the results on pass-through into U.S. import and export prices using our constructed unit value indices are wholly consistent with the estimates in Casas et al. (2016) and Gopinath and Rigobon (2008) that are based on BLS data. Lastly, Boz and Cerutti (2017) find favorable results when comparing country-level indices with those from the WTO and IMF World Economic Outlook.

Currency invoicing share. For currency invoicing shares we use the data set constructed by Gopinath (2015). The invoicing shares tend to be fairly stable over time so we take their simple averages over the years in which they are reported during 1999–2014. Appendix A.1 lists the USD and euro import invoicing share for the 39 countries in our sample with available invoicing data.

Country-level and global macro data. We use the World Bank's World Development Indicators (WDI) database as the source for annual average exchange rates and macroeconomic data for the world and our country sample. We obtain the WTI oil price, 1-year Treasury bill rate, and VIX from the St. Louis Fed's FRED database. See Appendix A.1 for details.

For some exercises below, we look at heterogeneity across advanced and emerging economies. We use the October 2017 IMF World Economic Outlook grouping of advanced economies, and label all other countries as emerging. This yields 31 advanced and 24 emerging economies listed in Appendix A.1.

4 Bilateral terms of trade, pass-through, and the dollar

In this section we show that, consistent with DCP, the U.S. dollar plays an outsized role in driving international trade prices and quantities. We first document that bilateral terms of trade are essentially uncorrelated with bilateral exchange rates. Next, we demonstrate that the bilateral (importer vs. exporter) exchange rates matter less than the exchange rate vis-à-vis the U.S. dollar for pass-through and trade elasticities of the average country in our sample. We find the euro to be much less important than the dollar. The effects of the dollar are stronger when the importing country has a higher fraction of trade invoiced in dollars. The dollar's role is greatest for trade between emerging market pairs, consistent with their higher reliance on dollar pricing. Finally, we show that the overall strength of the U.S. dollar is a key predictor of gross trade and producer/consumer price inflation in the rest of the world.

4.1 Terms of trade and exchange rates

We first relate bilateral terms of trade to bilateral exchange rates using panel regressions. In this subsection, a cross-sectional unit is defined to be an *unordered* country pair, so that both trade flows between two countries i and j are associated with the cross-sectional unit $\{i, j\}$. Define p_{ij} to be the log price of goods exported from country i to country j measured in currency j and e_{ij} to be the log bilateral exchange rate between country i and country j expressed as the price of currency i in terms of currency j. Define the bilateral log terms of trade $tot_{ij} = p_{ij} - p_{ji} - e_{ij}$ (i.e., export and import price indices are measured in the same currency). Moreover, let ppi_{ij} denote the log ratio of the producer price index (PPI) in country i divided by PPI in country j, with indices expressed in the same currency.

We consider the following regressions:

$$\Delta tot_{ij,t} = \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \alpha_k \Delta e_{ij,t-k} + \varepsilon_{ij,t}, \tag{1}$$

$$\Delta tot_{ij,t} = \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \alpha_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \beta_k \Delta ppi_{ij,t-k} + \varepsilon_{ij,t}, \tag{2}$$

where λ_{ij} and δ_t are dyadic and time fixed effects. Regression Eq. (1) relates the growth rate of the bilateral terms of trade to the growth rate of the bilateral nominal exchange rate (and two lags). As discussed in Section 2, if exporting firms set prices in their local currencies as in PCP and prices are sticky, the contemporaneous exchange rate coefficient α_0 should equal 1. On the other

Terms of trade and exchange rates

	unweighted		trade-weighted	
	(1)	(2)	(3)	(4)
VARIABLES	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$
$\Delta e_{ij,t}$	0.0369***	-0.00938	0.0813***	0.0218
	(0.00863)	(0.0130)	(0.0235)	(0.0317)
ΔER lags	2	2	2	2
PPI	no	yes	no	yes
Time FE	yes	yes	yes	yes
R-squared	0.008	0.011	0.028	0.042
Observations	24,270	19,847	24,270	19,847
Dyads	1,347	1,200	1,347	1,200

Table 1: The first (resp., last) two columns use unweighted (resp. trade-weighted) regressions. S.e. clustered by dyad. The number of dyads is about half that in Table 3 since here the two ordered country tuples (i, j) and (j, i) are collapsed into one cross-sectional unit $\{i, j\}$. *** p<0.01, ** p<0.05, * p<0.1.

hand, if exporting firms set prices in the destination currency as in LCP and prices are sticky, the contemporaneous exchange rate coefficient should be -1. However, if most prices are invoiced in U.S. dollars and are sticky in nominal terms, the coefficients α_k should be close to zero. The regression specified in Eq. (2) controls for lags 0–2 of the growth rate of the ratio of PPI in both countries, since firms' optimal reset prices should fluctuate with domestic cost conditions.

We consider both unweighted and trade-weighted regressions. To obtain trade weights, for each dyad and year, we compute the share of world non-commodities trade value (in USD) attributable to that dyad. Then, for each dyad, we compute the average share across the years 1992–2015.

In line with DCP, we find that bilateral exchange rates are virtually uncorrelated with bilateral terms of trade. The results of the panel regressions are shown in Table 1. If we do not control for relative PPI, the regression results indicate that the contemporaneous effect of the exchange rate on the terms of trade is positive. While the sign is consistent with PCP, the magnitude is not, as the 95% confidence interval equals $[0.02, 0.05]$ in the unweighted regression, and $[0.04, 0.13]$ in the weighted regression.[8] The coefficients on the lags (not reported) are also small in magnitude. When controlling for relative PPI, the point estimates of the coefficients on the bilateral exchange

[8]Attenuation bias is not a worry in this context, since the explanatory variables of interest (exchange rates) are precisely measured, except perhaps for time aggregation issues at the annual frequency.

Terms of trade and exchange rates: Country group heterogeneity

	unweighted			trade-weighted		
	(1)	(2)	(3)	(4)	(5)	(6)
	E↔E	E↔A	A↔A	E↔E	E↔A	A↔A
VARIABLES	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$	$\Delta tot_{ij,t}$
$\Delta e_{ij,t}$	0.0189	0.0480***	0.0182	0.0508***	0.111***	0.0220
	(0.0173)	(0.0110)	(0.0256)	(0.0176)	(0.0310)	(0.0473)
ΔER lags	2	2	2	2	2	2
PPI	no	no	no	no	no	no
Time FE	yes	yes	yes	yes	yes	yes
R-squared	0.028	0.011	0.008	0.051	0.078	0.025
Observations	3,527	11,857	8,886	3,527	11,857	8,886
Dyads	217	670	460	217	670	460

Table 2: "E↔A", say, denotes goods flows between Emerging and Advanced economies. The first (resp., last) three columns use unweighted (resp. trade-weighted) regressions, as in specifications (1) and (3) of Table 1. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

rate shrink further toward zero, and confidence intervals remain narrow. Hence, our results lend strong support to DCP: the terms of trade are unresponsive to bilateral exchange rates. Although the lack of correlation could in principle be consistent with a world of 50% PCP and 50% LCP, the next subsections refute that possibility. In addition, while the lack of correlation is consistent with any currency being a dominant currency, we provide evidence next that the major dominant currency is indeed the dollar. Lastly, the stability of the terms of trade for the average country in our sample cannot be explained by a model with flexible prices and strategic complementarities in pricing as in Atkeson and Burstein (2008) and Itskhoki and Mukhin (2017) because, as we show next, the import pass-through into destination country prices at-the-dock is high, contrary to the presence of strong complementarities in pricing. Further, as strategic complementarities get very large and pass-through at the dock low, the model behaves like LCP where the terms of trade becomes strongly negatively correlated with the exchange rate.[9]

Table 2 demonstrates that the terms of trade are nearly uncorrelated with the bilateral exchange rate across all advanced/emerging economy trade flows. We consider three subsamples, according to which zero, one, or both trading partners are advanced economies. The (fixed effect)

[9]Consistent with this observation, Table 2 in Itskhoki and Mukhin (2017) shows that their model generates correlations of terms of trade and exchange rates that are either 1 or -1, very different from the estimates in the data.

correlation of the terms of trade with the bilateral exchange rate is estimated to be positive in all cases, although not always statistically significantly so. The regression coefficient is the largest for trading relationships involving one country from each group, but even then the coefficient is merely 0.05 in unweighted regressions and 0.11 in trade-weighted regressions.

4.2 Exchange rate pass-through into prices

Exchange rate pass-through regressions are reduced-form regressions that relate price changes to exchange rate changes and other control variables relevant for pricing. We follow the literature and estimate the standard pass-through regression as described in Burstein and Gopinath (2014). Specifically, we estimate

$$\Delta p_{ij,t} = \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \alpha_k \Delta e_{ij,t-k} + \theta' X_{i,t} + \varepsilon_{ij,t}, \tag{3}$$

where λ_{ij} and δ_t are dyadic and time fixed effects. $X_{i,t}$ are other controls, namely the change in the log producer price index of the exporting country i measured in currency i (and two lags).[10] In this subsection and henceforth, the cross-sectional unit is an *ordered* country pair (i, j).

We modify this standard regression by including the dollar exchange rate, i.e., the log price $e_{\$j}$ of a U.S. dollar in currency j, alongside the bilateral exchange rate, as suggested in Section 2:

$$\Delta p_{ij,t} = \lambda_{ij} + \delta_t + \sum_{k=0}^{2} \alpha_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \beta_k \Delta e_{\$j,t-k} + \theta' X_{i,t} + \varepsilon_{ij,t}. \tag{4}$$

Lastly, we interact the dollar exchange rate with the importing country's dollar invoicing share:

$$\begin{aligned} \Delta p_{ij,t} = \lambda_{ij} + \delta_t &+ \sum_{k=0}^{2} \alpha_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \eta_k \Delta e_{ij,t-k} \times S_j \\ &+ \sum_{k=0}^{2} \beta_k \Delta e_{\$j,t-k} + \sum_{k=0}^{2} \psi_k \Delta e_{\$j,t-k} \times S_j + \theta' X_{i,t} + \varepsilon_{ij,t}. \end{aligned} \tag{5}$$

The estimates from Eq. (3) are reported in columns (1) and (4) of Table 3, corresponding to unweighted and trade-weighted regressions, respectively.[11] According to the regression estimates,

[10] Appendix B.2.2 shows that our results are robust to adding importer PPI and GDP growth as additional control variables.

[11] Henceforth, the trade weights are given by the average (across the years 1992–2015) share of world non-commodities trade value attributable to an *ordered* dyad (i, j).

Exchange rate pass-through into prices

VARIABLES	unweighted (1) $\Delta p_{ij,t}$	unweighted (2) $\Delta p_{ij,t}$	unweighted (3) $\Delta p_{ij,t}$	trade-weighted (4) $\Delta p_{ij,t}$	trade-weighted (5) $\Delta p_{ij,t}$	trade-weighted (6) $\Delta p_{ij,t}$
$\Delta e_{ij,t}$	0.757***	0.164***	0.209***	0.765***	0.345***	0.445***
	(0.0132)	(0.0126)	(0.0169)	(0.0395)	(0.0449)	(0.0336)
$\Delta e_{ij,t} \times S_j$			-0.0841***			-0.253***
			(0.0240)			(0.0482)
$\Delta e_{\$j,t}$		0.781***	0.565***		0.582***	0.120*
		(0.0143)	(0.0283)		(0.0377)	(0.0622)
$\Delta e_{\$j,t} \times S_j$			0.348***			0.756***
			(0.0326)			(0.0796)
ΔER lags	2	2	2	2	2	2
Exp. PPI	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes
R-squared	0.356	0.398	0.515	0.339	0.371	0.644
Observations	46,820	46,820	34,513	46,820	46,820	34,513
Dyads	2,647	2,647	1,900	2,647	2,647	1,900

Table 3: The first (resp., last) three columns use unweighted (resp. trade-weighted) regressions. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

when country j's currency depreciates relative to country i by 10%, import prices in country j rise by 8%, suggestive of close to complete pass-through at the one year horizon.[12] The second and third lags (not reported) are economically less important.

Columns (2) and (5) report estimates from regression Eq. (4). Including the dollar exchange rate sharply reduces the relevance of the bilateral exchange rate. It knocks the coefficient on the bilateral exchange rate from 0.76 to 0.16 in the unweighted regression, and from 0.77 to 0.34 in the weighted regression. Instead, almost all of the effect is absorbed by the dollar exchange rate.[13] Notice that, due to our inclusion of time fixed effects, the apparent dominance of the dollar cannot be an artifact of special conditions that may apply in times when the dollar appreciates or

[12] With year fixed effects this should be interpreted as fluctuations in excess of world annual fluctuations.

[13] In the literature, *unilateral* exchange rate pass-through is sometimes estimated using a Vector Error Correction Model (VECM) that allows for cointegration between price levels and exchange rates. However, Burstein and Gopinath (2014, p. 403) find VECM results to be highly unstable across specifications, and this issue is likely to be compounded by measurement error in our bilateral data.

depreciates against *all* other currencies, for example due to global recessions or flight to safety in asset markets. Appendix B.2.2 shows that our results are robust to the choice of time sample, including removing the post-2008 period.

The cross-dyad heterogeneity in pass-through coefficients is related to the propensity to invoice imports in dollars. Columns (3) and (6) interact the dollar and bilateral exchange rates with the share of invoicing in dollars at the importer country level, as in regression Eq. (5). Notice that we do not have data on the fraction of *bilateral* trade invoiced in dollars, so we use the importer's country-level share as a proxy. As expected, the import invoicing share plays an economically and statistically significant role for the dollar pass-through. Depending on whether we use trade weights or not, the regression results indicate that increasing the dollar invoicing share by 10 percentage points causes the contemporaneous dollar pass-through to increase by 3.5–7.6 percentage points. The R^2 values of the panel regressions are substantially improved by adding the invoicing share interaction terms. We further quantify the importance of the dollar invoicing share for explaining the cross-sectional variation in pass-through in Section 5.

Figs. 1 and 2 depict the regression results visually in the form of impulse response functions. Fig. 1 shows the impulse responses of the bilateral import price *level*. The top row shows unweighted regression results, the bottom row uses trade weights as described above. The left column shows the bilateral pass-through in the specifications without the dollar exchange rate, while the right column compares the bilateral and dollar pass-throughs in specifications with both exchange rates. Fig. 2 illustrates the pass-through heterogeneity as a function of the invoicing share S_j, as implied by the regression specifications with interactions. The figure focuses on three dollar shares: $S_j = 0.13$ (corresponding to Switzerland), $S_j = 0.59$ (Turkey), and $S_j = 0.88$ (Argentina). As depicted in Fig. 2, dollar pass-through is highest for Argentina with the largest dollar invoicing share and the least for Switzerland with its low dollar share. In the trade-weighted regressions (bottom row of Fig. 2), dollar pass-through is lower than bilateral pass-through for Switzerland, and that ranking is flipped for the case of Turkey and Argentina.

Table 4 shows that dollar dominance holds up qualitatively across flows between different country groups. We estimate Eq. (4) separately on the four subsamples corresponding to whether the exporter or importer is advanced or emerging. The contemporaneous dollar pass-through is larger than the bilateral pass-through in every subsample, except for trade flows between advanced economies estimated using trade weights. Even in this case, the dollar pass-through is close to the bilateral pass-through. The bilateral pass-through is particularly small for flows where the exporting country is emerging, and our conclusion of the dominance of the dollar is strongest for flows between emerging markets. These facts are in line with the results in Table 3 for regres-

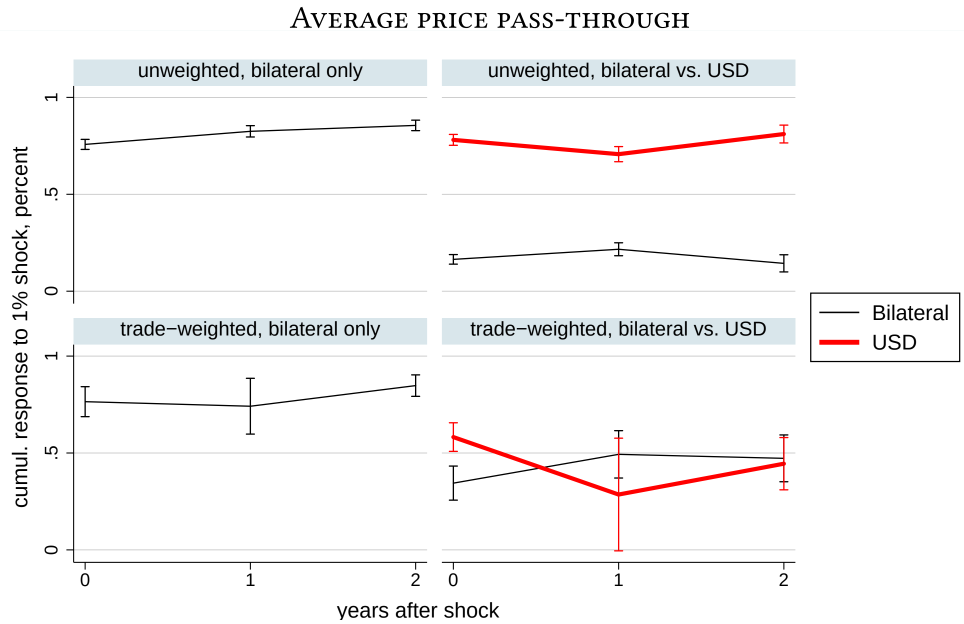


Figure 1: Impulse responses of bilateral price level to bilateral $e_{ij,t}$ and USD $e_{\$j,t}$ exchange rates. Based on the regressions in Table 3 without interactions. Top row: unweighted regression, bottom row: trade-weighted. Left column: specifications (1) and (4), right column: specifications (2) and (5). Error bars: 95% confidence intervals, clustering by dyad.

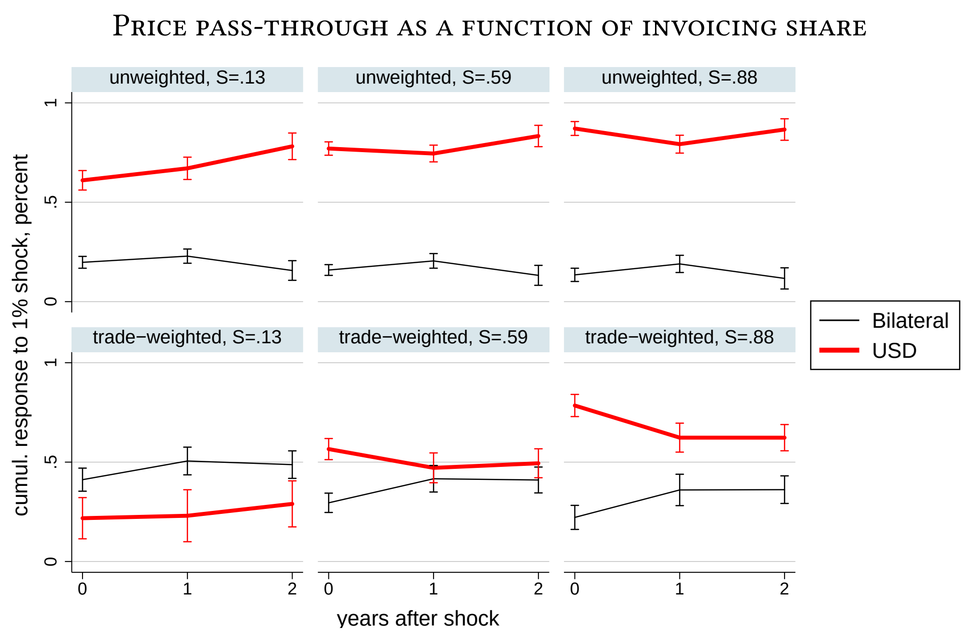


Figure 2: Impulse responses of bilateral price level to bilateral $e_{ij,t}$ and USD $e_{\$j,t}$ exchange rates, as a function of importer's dollar invoicing share S_j. Based on regression specifications (3) and (6) in Table 3 with interactions. Top row: unweighted regression, bottom row: trade-weighted. Error bars: 95% confidence intervals, clustering by dyad.

Exchange rate pass-through into prices: Country group heterogeneity

VARIABLES	unweighted (1) E→E $\Delta p_{ij,t}$	(2) E→A $\Delta p_{ij,t}$	(3) A→E $\Delta p_{ij,t}$	(4) A→A $\Delta p_{ij,t}$	trade-weighted (5) E→E $\Delta p_{ij,t}$	(6) E→A $\Delta p_{ij,t}$	(7) A→E $\Delta p_{ij,t}$	(8) A→A $\Delta p_{ij,t}$
$\Delta e_{ij,t}$	0.0980***	0.0514**	0.265***	0.332***	0.150***	0.150***	0.433***	0.373***
	(0.0329)	(0.0225)	(0.0379)	(0.0195)	(0.0391)	(0.0269)	(0.132)	(0.0504)
$\Delta e_{\$j,t}$	0.858***	0.766***	0.710***	0.409***	0.820***	0.498***	0.608***	0.287***
	(0.0353)	(0.0364)	(0.0382)	(0.0284)	(0.0487)	(0.0533)	(0.122)	(0.0487)
ΔER lags	2	2	2	2	2	2	2	2
Exp. PPI	yes	yes	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes	yes	yes
R-squared	0.470	0.152	0.530	0.142	0.572	0.252	0.467	0.264
Observations	6,763	10,589	12,318	17,150	6,763	10,589	12,318	17,150
Dyads	435	618	700	894	435	618	700	894

Table 4: "E→A", say, denotes goods flows from Emerging to Advanced economies. The first (resp., last) four columns use unweighted (resp. trade-weighted) regressions as in specifications (2) and (5) of Table 3. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

sions that interact with the dollar invoicing share, since emerging markets tend to have higher dollar invoicing shares. However, we stress that the importance of the dollar exchange rate is not limited to flows involving emerging markets.

4.3 Trade volume elasticity

Having demonstrated the outsized role of the U.S. dollar in determining international prices, we now show that the dollar also dominates the bilateral exchange rate when predicting bilateral trade volumes.

Table 5 shows the results from panel regressions of trade volumes on bilateral and dollar exchange rates. Let y_{ij} denote the log volume of goods exported from country i to country j. Our volume regressions take the same form as in the price pass-through regressions, Eqs. (3) to (5), except that the dependent variable is now the log growth rate $\Delta y_{ij,t}$ of bilateral trade volumes, and the extra controls $X_{j,t}$ (here indexed by j rather than i) consist of the log growth rate of real GDP (and two lags) for the importing country j. These regressions do not capture structural demand elasticity parameters, since we do not attempt to control for all relevant relative prices, and the

Trade elasticity with respect to exchange rate

	unweighted			trade-weighted		
	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$
$\Delta e_{ij,t}$	-0.119***	-0.0310*	-0.0765*	-0.0901***	-0.0163	-0.0971**
	(0.0139)	(0.0160)	(0.0403)	(0.0182)	(0.0236)	(0.0380)
$\Delta e_{ij,t} \times S_j$			0.118*			0.124**
			(0.0684)			(0.0519)
$\Delta e_{\$j,t}$		-0.186***	-0.140**		-0.155***	-0.131**
		(0.0250)	(0.0600)		(0.0277)	(0.0658)
$\Delta e_{\$j,t} \times S_j$			-0.0903			-0.00581
			(0.0871)			(0.0846)
ΔER lags	2	2	2	2	2	2
Imp. GDP	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes
R-squared	0.069	0.071	0.074	0.172	0.179	0.215
Observations	52,272	52,272	38,582	52,272	52,272	38,582
Dyads	2,807	2,807	2,014	2,807	2,807	2,014

Table 5: The first (resp., last) three columns use unweighted (resp. trade-weighted) regressions. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

importer's GDP growth is an imperfect proxy for the level of import demand. In particular, our results will invariably conflate expenditure switching and shifts in aggregate import demand. We view these regressions as predictive relationships that may inform future structural estimation exercises. Nevertheless, we will refer to the coefficients on exchange rates as "trade elasticities" for simplicity.

The volume regressions underline the dominant role played by the U.S. dollar. As in the case of the price pass-through regressions, adding the dollar exchange rate to the volume regressions knocks down the coefficient on the bilateral exchange rate by a substantial amount. The contemporaneous elasticity for the dollar exchange rate is about -0.19 to -0.13 across specifications, while the elasticity for the bilateral exchange rate is an order of magnitude smaller. Unlike the price pass-through regressions, the interactions of exchange rate changes with the importer's dollar invoicing share are mostly imprecisely estimated here.

Fig. 3 visually depicts the regression results in the form of impulse responses. The figure shows

Average trade elasticity

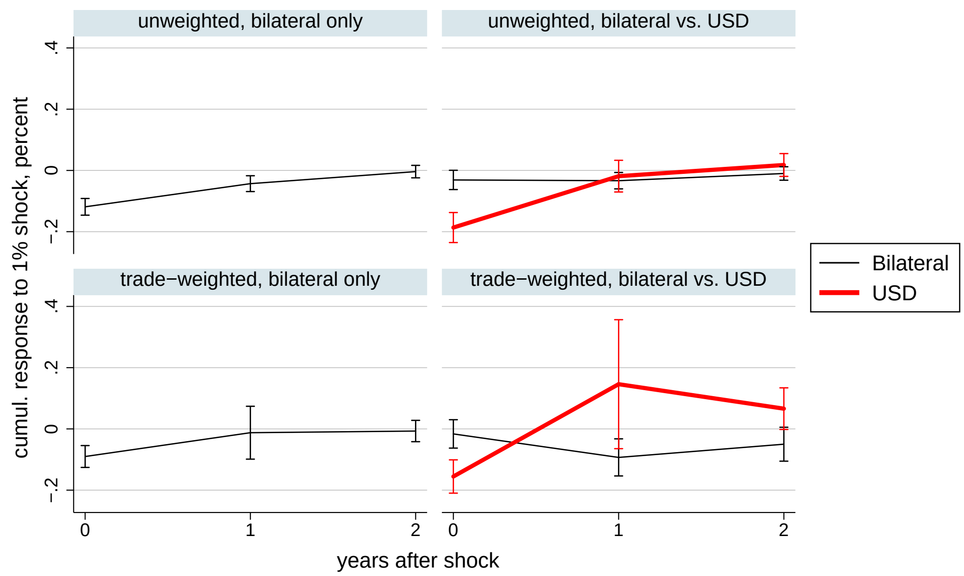


Figure 3: Impulse responses of bilateral volume to bilateral $e_{ij,t}$ and USD $e_{\$j,t}$ exchange rates. Based on regressions in Table 5 without interactions. Top row: unweighted regression, bottom row: trade-weighted. Left column: specifications (1) and (4), right column: specifications (2) and (5). Error bars: 95% confidence intervals, clustering by dyad.

the response of the level of bilateral trade volume to exchange rate shocks. The right column shows results from regressions with both bilateral and dollar exchange rates. It is apparent from the figure that the dollar exchange rate has a much more negative impact effect than the bilateral exchange rate. Yet, the figure also shows that the effect of either exchange rate on the level is essentially neutral at horizons of 1–2 years. One potential explanation is that the ratio of import prices and domestic prices adjust with a lag to exchange rate changes, implying that a year after the initial shock, relative prices faced by consumers are mostly unchanged compared to the period before the shock. However, we show in Appendix B.2.2 that this particular finding is driven by the early years in our sample, as results on the 2002–2015 subsample point toward a large and persistent negative effect of dollar appreciations on the volume of bilateral trade.

Table 6 shows that the contemporaneous trade elasticity of the dollar dominates the bilateral exchange rate elasticity in most breakdowns of emerging/advanced economy trade flows. It is only in flows from advanced to emerging markets that the bilateral exchange rate elasticity is estimated to be larger in absolute value than the dollar elasticity (and only in the unweighted regression), but the standard errors are particularly large in this subsample. On balance, the data indicates that a substantially negative dollar elasticity, coupled with a smaller bilateral exchange rate elasticity, is a common feature to emerging and advanced economy trade flows.

Trade elasticity with respect to exchange rate: Country group heterogeneity

	unweighted				trade-weighted			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	E→E	E→A	A→E	A→A	E→E	E→A	A→E	A→A
VARIABLES	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$
$\Delta e_{ij,t}$	-0.0488	-0.0145	-0.182***	-0.0737	-0.0471	-0.0441**	-0.0377	0.0228
	(0.0333)	(0.0212)	(0.0700)	(0.0481)	(0.0357)	(0.0225)	(0.117)	(0.0518)
$\Delta e_{\$j,t}$	-0.163***	-0.435***	0.00868	-0.340***	-0.208***	-0.251***	-0.0995	-0.302***
	(0.0588)	(0.0749)	(0.0704)	(0.0607)	(0.0641)	(0.0622)	(0.118)	(0.0548)
ΔER lags	2	2	2	2	2	2	2	2
Imp. GDP	yes	yes	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes	yes	yes
R-squared	0.093	0.049	0.100	0.082	0.237	0.301	0.218	0.214
Observations	8,239	12,967	12,932	18,134	8,239	12,967	12,932	18,134
Dyads	485	679	719	924	485	679	719	924

Table 6: "E→A", say, denotes goods flows from Emerging to Advanced economies. The first (resp., last) four columns use unweighted (resp. trade-weighted) regressions as in specifications (2) and (5) of Table 5. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

4.4 Trade flows to and from the U.S.

The data is consistent with an additional key prediction of DCP: Trade flows involving the U.S. as a trading partner are special. Specifically, in this section we show that bilateral exchange rate pass-through into U.S. exports is complete and immediate, while U.S. import prices are insensitive to bilateral exchange rates. Moreover, U.S. import volumes are insensitive to the bilateral exchange rate, as predicted by theory.

Fig. 4 shows impulse response functions of import prices for goods flowing from or to the U.S. These figures are obtained from panel regressions as in the baseline unweighted fixed effects specifications in Sections 4.2 and 4.3, except we drop the dollar exchange rate (but preserve the time fixed effects). Consistent with the very high fraction of U.S. exports and imports being invoiced in dollars (97% and 93%, respectively), bilateral exchange rate pass-through into prices is 100% on impact for U.S. exports and close to zero for U.S. imports.

Table 7 confirms that U.S. import volumes are insensitive to bilateral exchange rates, unlike the imports of the rest of the world from the U.S. We run a fixed effects regression of trade volume growth on lagged bilateral exchange rates, importer GDP, and year fixed effects, as in

Price pass-through: Flows to and from U.S.

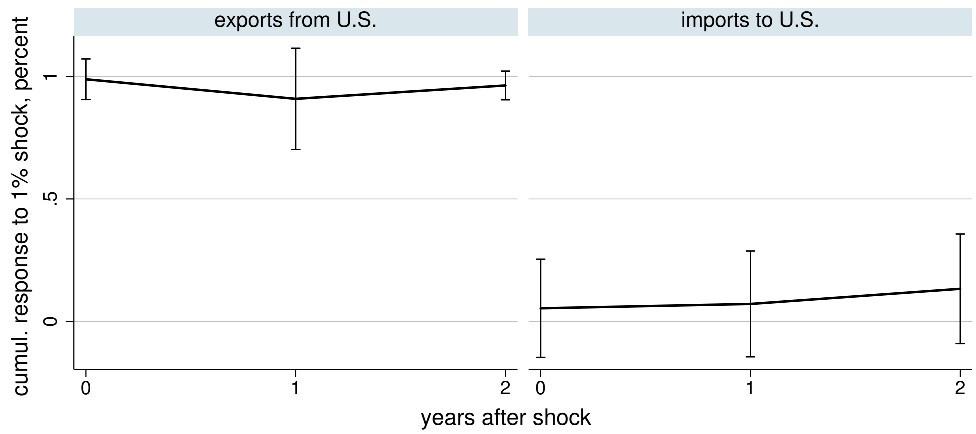


Figure 4: Impulse response of bilateral price level to bilateral exchange rate $e_{ij,t}$. Left column: U.S. exports, right column: U.S. imports. Error bars: 95% confidence intervals, clustering by dyad and applying small-sample "LZ2-BM" adjustment in Imbens and Kolesár (2016).

Trade elasticity: U.S. vs. non-U.S. imports

	unweighted	trade-weighted
	(1)	(2)
VARIABLES	$\Delta y_{ij,t}$	$\Delta y_{ij,t}$
$\Delta e_{ij,t}$	-0.121***	-0.107***
	(0.0141)	(0.0194)
$\Delta e_{ij,t} \times$ ImpUS	0.124***	0.117***
	(0.0329)	(0.0318)
ΔER lags	2	2
Imp. GDP x ImpUS	yes	yes
Time x ImpUS FE	yes	yes
R-squared	0.069	0.180
Observations	52,272	52,272
Dyads	2,807	2,807

Table 7: "ImpUS" is in indicator for whether importing country is the U.S. S.e. clustered by dyad. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

specifications (1) and (4) of Table 5. Here, however, we additionally interact all right-hand side variables with an indicator for whether the importing country is the U.S. When the importing country is *not* the U.S., the within-year bilateral trade volume response is estimated at -0.12% (unweighted) following a 1% depreciation of the importer currency, similar to the all-country regression in Table 5. In contrast, we find U.S. imports to be completely insensitive to the bilateral exchange rate on impact, with an implied contemporaneous import volume response of 0.003% following a 1% depreciation of the dollar. The difference between the contemporaneous import elasticity for the U.S. vs. that for the rest of the world is highly significant. Hence, the data indicates that U.S. trade balance adjustment following exchange rate movements occurs primarily through exports rather than imports, a consequence of the predominance of dollar invoicing in U.S. trade.

4.5 The dollar versus the euro

We now compare the explanatory power of the dollar exchange rate with that of the euro. We show that the dollar dominates both the bilateral exchange rate and the euro in regression specifications that include all three exchange rates.

The preceding panel regressions do not directly imply that the U.S. dollar is a uniquely important vehicle currency. In our regression specifications *without* interactions, we would have obtained exactly the same coefficient estimates if we had used the euro exchange rate, say, in place of the dollar exchange rate, since we control for time fixed effects. Nevertheless, our specifications *with* interactions indicated that the dollar plays a special role. Now we directly compare the explanatory power of the dollar against that of the euro in panel regressions that do not control for time fixed effects but instead control for observed global real and financial variables.

To measure bilateral price pass-through from the dollar and the euro, we run panel regressions of the form

$$\Delta p_{ij,t} = \lambda_{ij} + \sum_{k=0}^{2} \alpha_k \Delta e_{ij,t-k} + \sum_{k=0}^{2} \beta_k \Delta e_{\$j,t-k} + \sum_{k=0}^{2} \xi_k \Delta e_{€j,t-k} + \theta' X_{ij,t} + \varepsilon_{ij,t}, \tag{6}$$

where $e_{€j,t}$ denotes the log euro exchange rate in units of currency j per euro. Notice that we omit time fixed effects, as is necessary to identify β_k and ξ_k separately. In addition to lags 0–2 of exporter PPI log growth, the controls $X_{ij,t}$ consist of the contemporaneous values of global real GDP growth, global GDP deflator inflation, global export volume growth, growth in the WTI oil price deflated by the global GDP deflator, and the log VIX. The time sample for regressions in this subsection is 2002–2015 due to the introduction of the euro in 1999 and our use of lagged exchange rate changes.

Fig. 5 shows that the euro pass-through into prices is negligible on average, while the dollar pass-through remains high when we control for the euro. The figure displays the regression results in the form of impulse responses of the bilateral price level; corresponding regression tables are available in Appendix B.2.2. The left column shows results for specifications that do not include the dollar exchange rate, i.e., restricting $\beta_k = 0$ in Eq. (6). The right column displays estimates of bilateral, dollar, and euro pass-through from regressions with all three exchange rates. In the latter specifications, the dollar pass-through is quantitatively much larger than the euro pass-through; indeed, the results are close to those of Section 4.2 where we did not consider the euro at all. The euro pass-through is quantitatively small at all horizons, and it is in fact estimated to be slightly negative. The difference between the dollar and euro pass-through is statistically significant at conventional significance levels. In Appendix B.2.2 we show through regressions with interactions that the importer's dollar and euro invoicing shares help explain the heterogeneity in pass-through, with the expected signs.

Similarly, the dollar exchange rate has the largest predictive power for trade volumes. We run panel regressions similar to Eq. (6), except with volume growth $\Delta y_{ij,t}$ on the left-hand side, and we replace exporter PPI with lags 0–2 of importer real GDP growth in the list of controls $X_{ij,t}$. Fig. 6 shows impulse responses of the level of bilateral trade volume to the bilateral, dollar, and euro exchange rates. The dollar exchange rate is the only one of the three that has a quantitatively large negative association with trade volumes.[14]

4.6 Effect of U.S. dollar on rest-of-world trade and inflation

Underscoring the quantitative significance of DCP, we argue that the dollar has substantial predictive power for aggregate trade among countries in the rest of the world. That is, the dollar is important for predicting global trade, even when excluding countries' direct trade with the U.S. Specifically, a 1% U.S. dollar appreciation against all other currencies in the world predicts a 0.6% decline within a year in the volume of total trade between countries in the rest of the world, holding constant various proxies for the global business cycle.

We measure the elasticity of rest-of-world trade volume to the dollar by aggregating up from our richest bilateral panel regression specification. This produces results that exploit our panel data set, unlike a simple annual time series regression of global trade on an effective dollar ex-

[14] The different long-run level effect of the dollar in Figs. 3 and 6 is due to the difference in time sample, as discussed in Appendix B.2.2.

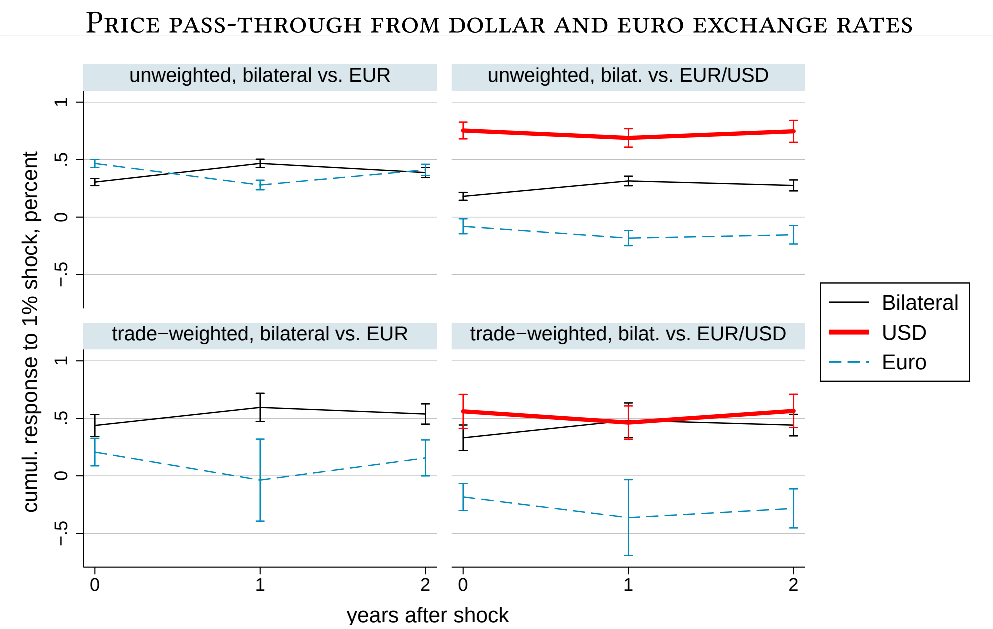


Figure 5: Impulse responses of bilateral price level to bilateral $e_{ij,t}$, USD $e_{\$j,t}$, and euro $e_{€j,t}$ exchange rates. Based on regressions in Table 11, Appendix B.2.2. Top row: unweighted regression, bottom row: trade-weighted. Left column: specifications with only bilateral and euro ER, right column: specifications adding USD. Error bars: 95% confidence intervals, clustering by dyad.

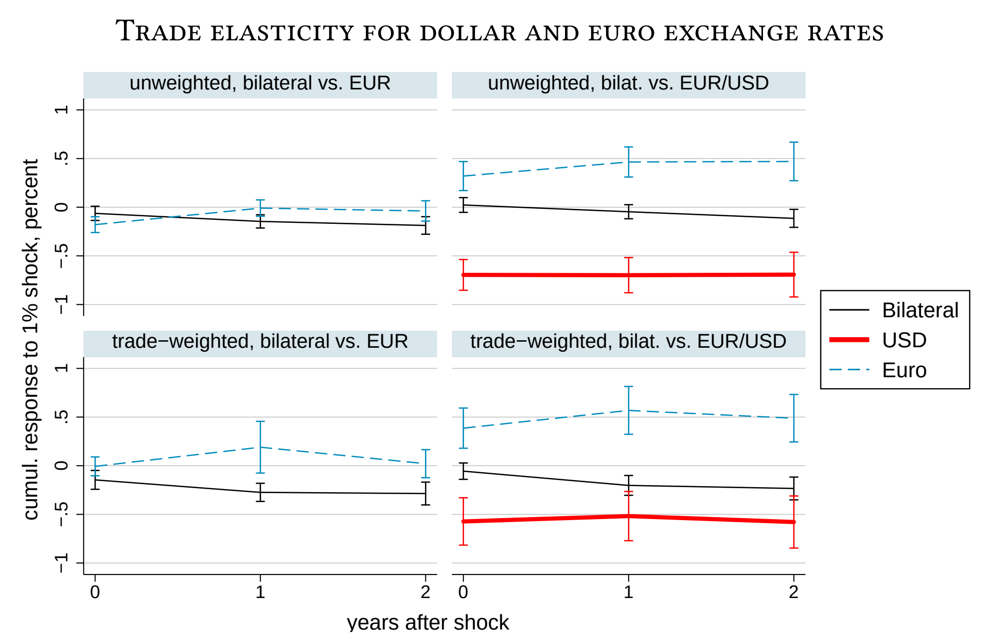


Figure 6: Impulse responses of bilateral volume to bilateral $e_{ij,t}$, USD $e_{\$j,t}$, and euro $e_{€j,t}$ exchange rates. Based on regressions in Table 12, Appendix B.2.2. Top row: unweighted regression, bottom row: trade-weighted. Left column: specifications with bilateral and euro ER, right column: specifications adding USD. Error bars: 95% confidence intervals, clustering by dyad.

change rate index. Consider the following regression model with bilateral, dollar, and euro exchange rates, as well as interactions with dollar and euro import invoicing shares:

$$\begin{aligned}\Delta y_{ij,t} = &\textstyle\sum_{k=0}^{2} \left(\alpha_k + \eta_k(1 - S_j - S_j^{€})\right)\Delta e_{ij,t-k} \\ &+ \textstyle\sum_{k=0}^{2} \left(\beta_k + \psi_k S_j\right)\Delta e_{\$j,t-k} \\ &+ \textstyle\sum_{k=0}^{2} \left(\xi_k + \vartheta_k S_j^{€}\right)\Delta e_{€j,t-k} \\ &+ \lambda_{ij} + \theta' X_{ij,t} + \varepsilon_{ij,t}.\end{aligned} \tag{7}$$

Here S_j and $S_j^{€}$ are the importer's country-level dollar and euro invoicing shares, respectively, and λ_{ij} is a dyad fixed effect. Because we are interested in the effect of a dollar appreciation against all other currencies, we do not control for time fixed effects. Instead, we control for the same proxies for the global business cycle as in Section 4.5, except world export volume growth. $X_{ij,t}$ also includes lags 0–2 of importer real GDP growth.

The main object of interest is the response of rest-of-world (i.e., ex-U.S.) aggregate trade volume to a 1% appreciation of the dollar relative to all other currencies, holding constant the global business cycle. Let w_j denote country j's total non-commodity import value from all countries except the U.S. in some reference year, normalized so that $\sum_{j\neq \text{US}} w_j = 1$.[15] We conceptualize the rest-of-world aggregate trade bundle as a Cobb-Douglas aggregate of individual-country (gross) imports with weights w_j. According to the bilateral interactive regression model (7), the *ceteris paribus* effect of a 1% dollar appreciation on $\sum_{j\neq \text{US}} w_j \Delta y_{ij,t}$, the weighted growth of rest-of-world imports from destination i, is given by

$$\textstyle\sum_{j\neq \text{US}} w_j(\beta_k + \psi_k S_j) = \beta_k + \psi_k \sum_{j\neq \text{US}} w_j S_j$$

k years after the appreciation, for each import destination i other than the U.S. Thus, to measure the response of rest-of-world aggregate imports to a dollar appreciation, we simply have to compute the impulse response of trade volume for an importer j whose U.S. dollar invoicing share happens to equal $\sum_{j\neq \text{US}} w_j S_j$, the weighted average dollar invoicing share, computed using our ex-U.S. import value weights w_j. In practice, w_j depends on the year in which import values are measured, but Appendix B.2.2 shows that the weighted average $\sum_{j\neq \text{US}} w_j S_j$ fluctuates little around a mean of 0.40 in the 2002–2015 sample, so we use the 0.40 value for our exercise.

Fig. 7 shows that rest-of-world aggregate import volume contracts markedly following an

[15]"All countries" refers to the world aggregate in Comtrade, not only the countries in our regression sample. Note that the weight w_j is different from the weights used in the trade-weighted regressions in Table 5 and elsewhere.

Response of rest-of-world aggregate trade to USD appreciation, 2002–2015

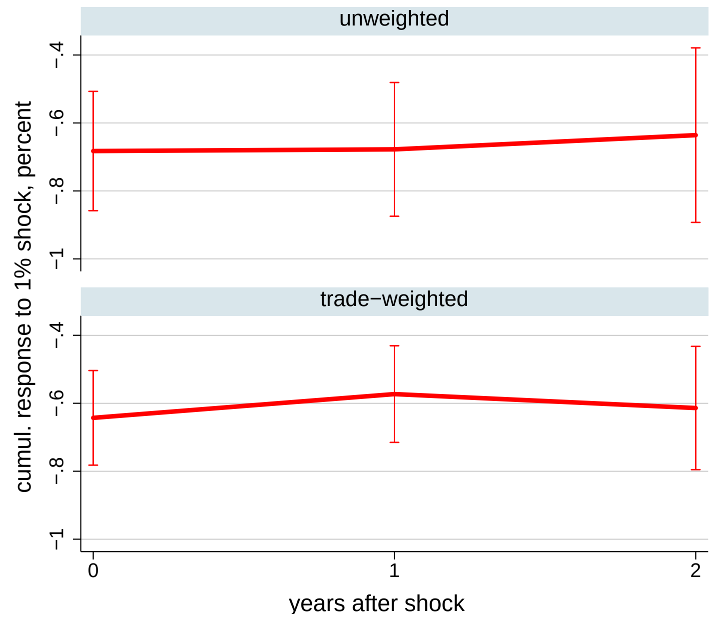


Figure 7: Impulse responses of rest-of-world aggregate trade volume to a 1% U.S. dollar appreciation against all other currencies, holding constant all other exchange rates and the global business cycle. Top row: unweighted regression, bottom row: trade-weighted. Error bars: 95% confidence intervals, clustering by dyad.

appreciation of the dollar against all other currencies. A 1% *ceteris paribus* dollar appreciation leads to a 0.6% contraction in rest-of-world trade volume within the year (regardless of whether we use unweighted or trade-weighted regressions), and this contractionary effect persists out to at least two years. Recall that the regression controls for various proxies for the global business and financial cycles. In unreported results, we find an even larger effect if we repeat the exercise on data for the pre-crisis period 2002–2007. While our regression specification cannot be interpreted structurally, the magnitude of the predictive effect underscores the importance of the dollar's role in world trade.

Finally, country-level regressions reveal significant dollar pass-through to foreign consumer and producer prices that increases with countries' dollar invoicing share in imports. Earlier work by Gopinath (2015) provides back-of-the-envelope calculations of dollar exchange rate spillovers on foreign consumer and producer prices based on estimated country-level import price pass-through and the import content of consumption. We take a more direct approach and regress countries' CPI or PPI on the dollar exchange rate as well as its interaction with the dollar invoicing share in imports using a specification with country and time fixed effects, detailed in Appendix B.2.1. The estimation is for the post-2002 (post-euro) sample as the full-sample results are influenced by a handful of countries with high inflation around large depreciation episodes of the 1990s. We find the average pass-through of the dollar into CPI (resp., PPI) to be 11% (resp., 28%) within the year. The specifications that include the dollar exchange rate interaction with the dollar invoicing share reveal that the dollar pass-through is higher for countries that have a larger fraction of their imports invoiced in dollars. The contemporaneous interaction term is statistically significant at the 10% level for both the CPI and PPI specifications, and also at the 5% level for the CPI specification. We caution, though, that the size of the pass-through is imprecisely estimated when controlling for country and time fixed effects.

5 Determinants of pass-through heterogeneity

This section shows that the cross-dyad variation in exchange rate price pass-through and trade elasticity is well explained by the dollar's dominance as invoicing currency. The theoretical framework underlying DCP predicts that pass-through from bilateral exchange rates to prices or quantities should vary across countries, depending on the share of imports invoiced in dollars. The panel regressions in the previous section indicate that this interaction effect is statistically and economically significant for price pass-through. In this section we quantify the interaction effect relative to unobserved factors affecting the cross-sectional heterogeneity of price pass-through and trade elasticities.

5.1 Bayesian model of pass-through heterogeneity

We employ a Bayesian hierarchical panel data model with cross-sectionally varying slopes. This model optimally exploits the geographical and temporal richness of our data set. By explicitly modeling the cross-sectional heterogeneity of pass-through, we are able to quantify how much of this heterogeneity can be explained by the share of trade invoiced in dollars (for brevity, here we use the term "pass-through" to describe the relationship between exchange rates and prices *or* quantities). Such questions cannot be answered by linear panel models with interactions, as these common-coefficients models are unable to quantify the *overall* cross-sectional heterogeneity of pass-through. Thus, we use a hierarchical Bayes framework with a nonparametric specification for the distribution of pass-through coefficients conditional on the dollar invoicing share.

The hierarchical approach lets the data determine the degree of variation in pass-through across trade dyads.[16] This approach can roughly be thought of as striking a balance between two

[16] At an abstract level, hierarchical Bayes methods treat certain prior parameters as unknown model parameters,

extreme but standard econometric methods. In one extreme, dyad-by-dyad time series regressions are run to determine dyad-specific pass-through coefficients. However, these pass-through estimates would be highly noisy due to the availability of on average about 20 annual data points per dyad, especially given the need to control for other covariates. In the other extreme, we could run constant-coefficient panel regressions as in Section 4, which are informative about average pass-through as well as interaction terms, but they are useless for estimating the extent and nature of the overall cross-sectional heterogeneity of pass-through. Our hierarchical Bayes approach models this heterogeneity directly and flexibly, allowing the entire panel data set to inform the estimates of the distribution of pass-through as well as individual pass-through coefficients. Being a fully Bayesian method, uncertainty assessment and model selection are straightforward.

Model. The outcome equation of the model is a linear panel data model with dyad and time fixed effects, except that some of the coefficients are allowed to vary across dyads:

$$Y_{ij,t} = \lambda_{ij} + \delta_t + \gamma_{ij}' R_{ij,t} + \theta' X_{ij,t} + \varepsilon_{ij,t}. \tag{8}$$

In our applications, the outcome $Y_{ij,t}$ will be price or quantity in log growth, while the covariates $R_{ij,t}$ with cross-sectionally varying coefficients γ_{ij} will be the contemporaneous log growth rates of the bilateral and U.S. dollar exchange rates, $R_{ij,t} = (\Delta e_{ij,t}, \Delta e_{\$j,t})'$. The covariates $X_{ij,t}$ with cross-sectionally constant coefficients θ include lags of the exchange rates as well as the other exogenous controls used in Section 4.[17] We impose a standard random effects assumption on the dyad-specific effects, $\lambda_{ij} \sim N(\alpha, \tau^2)$ (i.i.d. across dyads), and assume Gaussian errors $\varepsilon_{ij,t} \sim N(0, \sigma^2)$ (i.i.d. across dyads and time).[18] We place independent diffuse half-Cauchy priors on τ and σ and independent diffuse Cauchy priors on the intercept α, the time fixed effects δ_t, and the cross-sectionally constant coefficients θ. See Appendix A.2.1 for details on the prior.

To economize on the number of parameters, we assume that the *sum* of the pass-through coefficients on the bilateral and dollar exchange rates is constant across dyads: $\gamma_{ij,1} + \gamma_{ij,2} = \bar{\gamma}$ for all (i, j). This restriction is motivated by the institutional fact that, in most countries in our

which themselves are endowed with prior distributions that get updated by the data. This approach is similar to "empirical Bayes" or classical "random effects" methods, which in effect estimate the prior distribution (here: the distribution of pass-through coefficients) from the data.

[17] All probability statements in this section are conditional on the covariates $R_{ij,t}$ and $X_{ij,t}$. In particular, we assume strict exogeneity of all covariates, and γ_{ij} is independent of $X_{ij,t}$ conditional on $R_{ij,t}$.

[18] In the panel regressions in Section 4 we do not find evidence of economically significant serial correlation in the idiosyncratic errors. Identification of the full distribution of random slopes in linear panel data models is only possible under *a priori* restrictions on the persistence of the idiosyncratic regressions errors (Chamberlain, 1992; Arellano and Bonhomme, 2012).

sample, trade that is not invoiced in dollars is invoiced in local currency, so dyads with high dollar pass-through should exhibit low bilateral pass-through, and vice versa. The restriction on the vector γ_{ij} implies that the outcome equation can be written as

$$Y_{ij,t} = \lambda_{ij} + \delta_t + \gamma_{ij,1}(\Delta e_{\$j,t} - \Delta e_{ij,t}) + \bar{\gamma}\Delta e_{ij,t} + \theta' X_{ij,t} + \varepsilon_{ij,t}. \tag{9}$$

This restricted outcome equation can be written in the general form (8), with γ_{ij} a scalar, $R_{ij,t} = \Delta e_{\$j,t} - \Delta e_{ij,t}$, and subsuming the term $\bar{\gamma}\Delta e_{ij,t}$ in the covariate terms $\theta' X_{ij,t}$. We assume this notation in what follows.

A key object in the model is the cross-sectional distribution of dollar pass-through conditional on the dollar invoicing share. We continue to denote the importer's observed dollar invoicing share by S_j. For maximal flexibility, we use a nonparametric specification of the conditional dollar pass-through distribution $\gamma_{ij} \mid S_j$, while letting the hyperparameters of the prior be updated by the data. Specifically, we follow Pati et al. (2013) and Liu (2017) and assume that, conditional on the importer's dollar invoicing share, the dollar pass-through coefficient is drawn from a Mixture of Gaussian Linear Regressions (MGLR):

$$(\gamma_{ij} \mid S_j) \sim \begin{cases} N(\mu_{0,1} + \mu_{1,1}S_j, \omega_1^2) & \text{with prob. } \pi_1(S_j), \\ N(\mu_{0,2} + \mu_{1,2}S_j, \omega_2^2) & \text{with prob. } \pi_2(S_j), \\ \vdots & \\ N(\mu_{0,K} + \mu_{1,K}S_j, \omega_K^2) & \text{with prob. } \pi_K(S_j), \end{cases}$$

independent across dyads (i, j). Thus, the dollar pass-through γ_{ij} is drawn from one of K normal distributions, each with possibly different mean and variance parameters. The priors on the hyperparameters $\mu_{0,k}$, $\mu_{1,k}$, and ω_k are described in Appendix A.2.1. The mixture probabilities $\pi_k(S_j)$ are allowed to depend flexibly on the dollar share. We adopt the "probit stick-breaking" specification of Pati et al. (2013),

$$\pi_k(s) = \begin{cases} \Phi(\zeta_k(s)) \prod_{j=1}^{k-1}(1 - \Phi(\zeta_j(s))) & \text{for } k = 1, \dots, K-1, \\ 1 - \sum_{j=1}^{K-1} \pi_j(s) & \text{for } k = K, \end{cases}, \quad s \in [0, 1],$$

where $\Phi(\cdot)$ is the standard normal CDF. As in Liu (2017), we place independent nonparametric Gaussian process priors on the functions $\zeta_k(\cdot)$ for $k = 1, \dots, K-1$. See Appendix A.2.1.

The nonparametric prior on the cross-sectionally varying dollar pass-through coefficients allows the data to speak flexibly about our key question of interest, the extent to which the dollar invoicing share can explain pass-through heterogeneity. MGLR priors, as defined above, can accommodate a wide variety of shapes of the conditional density of $\gamma_{ij} \mid S_j$, including heavy-tailed, skewed, and multimodal conditional distributions. Since the mixture probabilities $\pi_k(S_j)$ depend on S_j, the functional form of the conditional distribution is allowed to change as the dollar invoicing share S_j varies. In particular, we do not impose that the distribution of γ_{ij} shifts linearly with S_j.[19] Pati et al. (2013) show that, if $K = \infty$, MGLR priors yield posterior consistency in nonparametric conditional density estimation problems under weak assumptions. We instead allow the data to inform us about the choice of the number K of mixture components, using the Bayesian Leave-One-Out (LOO) cross-validation model selection criterion of Gelfand et al. (1992) and Vehtari et al. (2017), cf. Appendix A.2.2.

Posterior sampling. We use the Bayesian statistics software package Stan to draw from the posterior distribution of the model parameters (Stan Development Team, 2016). Stan produces samples from the posterior using the No U-Turn Sampler of Hoffman and Gelman (2014), a variant of the Markov Chain Monte Carlo (MCMC) procedure Hamiltonian Monte Carlo (Neal, 2011). Stan achieves robust and rapid mixing in our high-dimensional hierarchical model, without requiring priors to be conjugate. Appendix B.3.4 details the performance of the MCMC routine.

5.2 Results: price pass-through

We find that the importer's share of dollar invoicing explains a substantial fraction of the heterogeneity in dollar pass-through into prices, confirming a key channel in DCP. Below we summarize the most important features of the posterior distribution for our purposes, while Appendix B.3.2 provides additional details on other parameters.

Our empirical specification broadly follows Section 4. In terms of the general Bayesian model in Eq. (8), we set $Y_{ij,t} = \Delta p_{ij,t}$. As extra covariates in $X_{ij,t}$, we use the exporter's log PPI growth and one lag each of log PPI growth, bilateral exchange rate log growth, and dollar exchange rate log growth (second lags were found to be unimportant in Section 4). We do not use regression weights in this section. We remove a few dyads whose data have gaps in the middle of the sample. Since we require data on the importer's dollar invoicing share, our final sample consists of 1856 dyads for a total of 35,398 observations (average of 19.1 years per dyad).

Our preferred specification uses $K = 2$ mixture components for the conditional distribution of dollar pass-through coefficients given the dollar invoicing share. The LOO model selection criterion indicates strong support for $K \geq 2$ against $K = 1$, but the criterion is mostly flat

[19] It is only the distribution *conditional on a mixture component* k that is assumed to shift linearly.

for $K = 2, 3, \ldots, 8$. Because the posterior summaries below are virtually unchanged across these values of K, we prefer to show results for the more parsimonious model $K = 2$ here. Appendix B.3.3 provides results for the richer $K = 8$ specification.

Fig. 8 shows that a higher importer (country-level) dollar invoicing share is associated with a rightward shift in the cross-sectional density of dollar pass-through. The figure focuses on three invoicing shares: a low one (Switzerland), a medium one (Turkey), and a high one (Argentina). While the cross-sectional heterogeneity in pass-through is large, there is a noticeable overall rightward shift in dollar pass-through when going from a low-S_j country to a high-S_j country. Based on posterior median estimates, the mode of the γ_{ij} distribution shifts by about 0.10 when the dollar invoicing share increases from Switzerland to Argentina levels. This is a substantial shift when compared to the estimated cross-dyad interquartile range of γ_{ij} of 0.13 (see below).[20]

Fig. 9 plots the posterior conditional mean and standard deviation of the conditional distribution $\gamma_{ij} \mid S_j$ across all observed values of S_j. The figure confirms that the three conditional densities plotted in Fig. 8 are representative of the entire observed distribution of S_j values. Although not assumed *a priori* by our model, the conditional mean $E[\gamma_{ij} \mid S_j]$ appears to be approximately linear, with a slope that is broadly consistent with the linear model with interactions in Section 4. The conditional standard deviation appears to be fairly constant across S_j values, although the posterior uncertainty is large. However, the conditional distributions are heavy-tailed, as evidenced by the fact that the LOO criterion strongly prefers the $K = 2$ mixture model to the $K = 1$ model with normally distributed heterogeneity.

Fig. 10 provides further evidence that dollar pass-through is high on average but highly heterogeneous, and about 15% of the cross-dyad variance of dollar pass-through is explained by the importer's dollar invoicing share. The figure shows histograms of the posterior draws of the cross-dyad median and interquartile range (IQR) of γ_{ij} for the 1856 dyads in the sample. The median dollar pass-through is consistent with the panel regressions in Section 4 (median median 0.76), but there is substantial heterogeneity in pass-through across dyads (median IQR 0.13), a fact we would not have been able to establish using standard linear panel regressions. The figure also plots the histogram of posterior draws of the cross-sectional correlation coefficient of γ_{ij} and S_j, after winsorizing γ_{ij} by 5% in each tail to reduce the influence of outlier dyads. There is a clear positive correlation (median correlation 0.39), again demonstrating that dyads with high dollar pass-through also tend to have a high importer dollar invoicing share. By squaring the correla-

[20]Recall that our data set is limited to using *country-level* dollar invoicing shares for the importer, S_j, as opposed to the ideal of dyad-specific invoicing shares. We conjecture that the quantitative importance of the importer's country-level dollar invoicing share provides a lower bound on the importance of the (unobserved) dyad-level invoicing share.

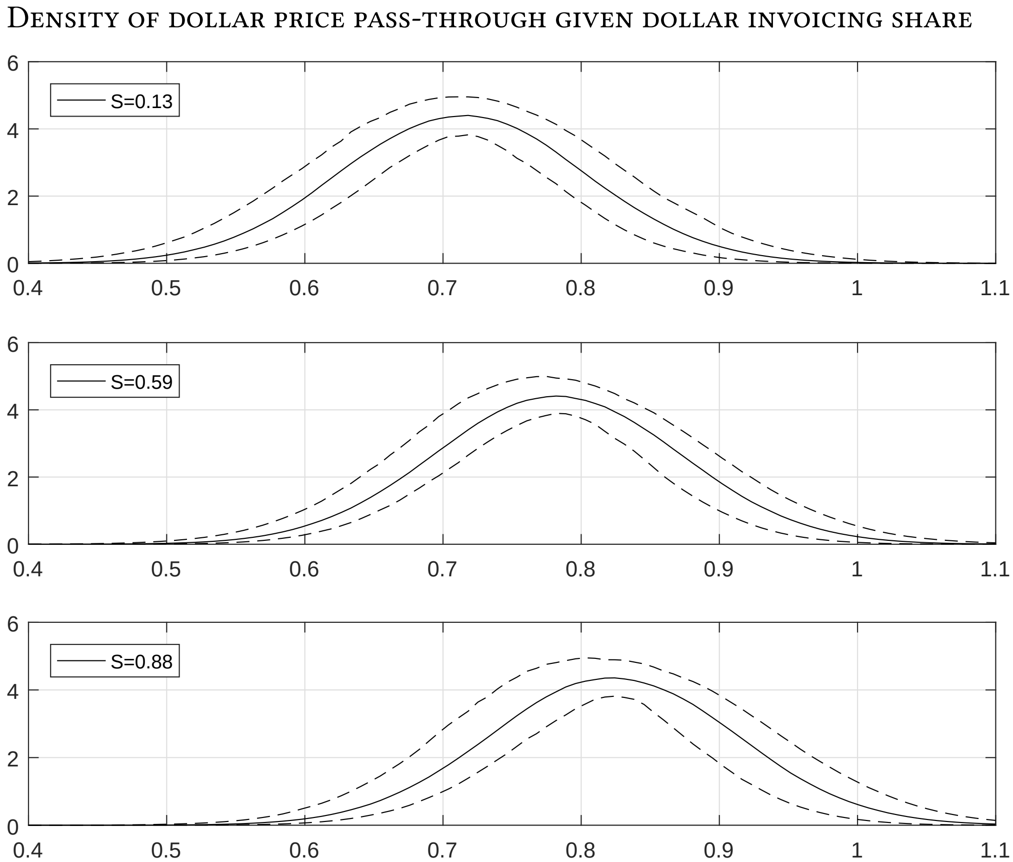


Figure 8: Model-implied conditional density $f(\gamma_{ij} \mid S_j)$ plotted at the dollar import invoicing shares S_j of Switzerland (top), Turkey (middle), and Argentina (bottom). Solid lines are posterior medians, dashed lines are 95% pointwise equal-tailed posterior credible intervals.

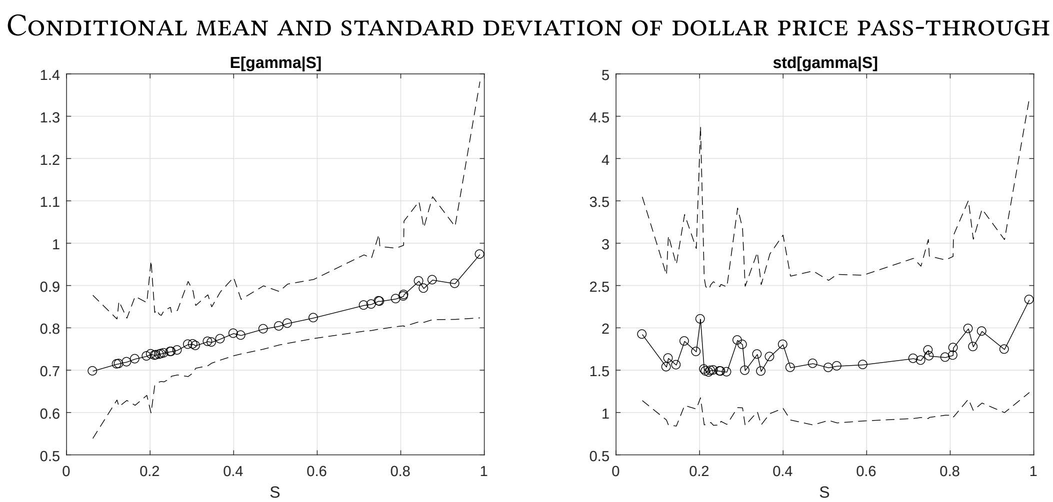


Figure 9: Model-implied conditional mean (left) and standard deviation (right) of γ_{ij} given S_j. Solid lines are posterior medians, dashed lines are 95% pointwise equal-tailed posterior credible intervals. Circles indicate observed S_j values.

Sample distribution of dollar price pass-through

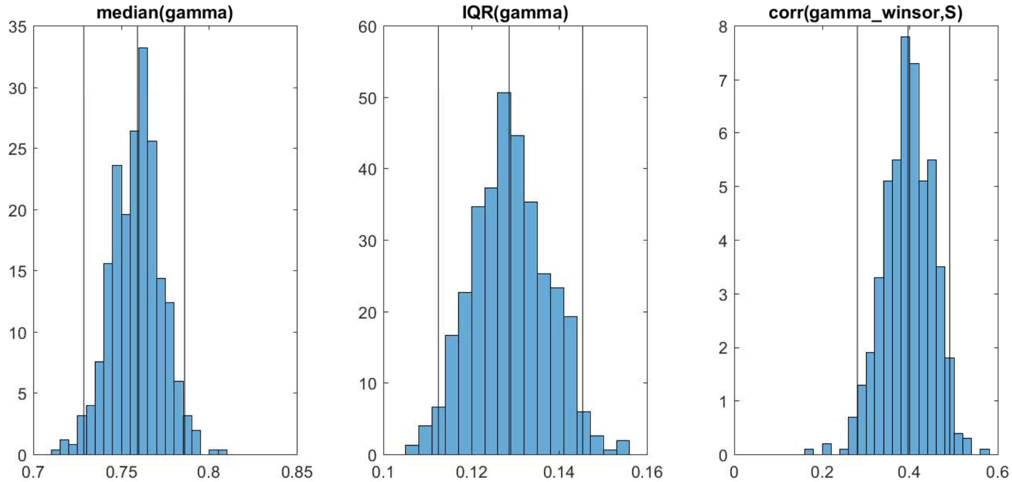


Figure 10: Histogram of posterior draws of the sample median of γ_{ij} (left), the sample interquartile range of γ_{ij} (middle), and winsorized correlation of γ_{ij} and S_j (right). That is, for each posterior draw, we compute the sample median, IQR, and winsorized correlation across the 1856 dyads in our sample. Vertical lines mark the 2.5, 50, and 97.5 posterior percentiles.

tion, we obtain the R^2 value in a cross-dyad regression of (winsorized) dollar pass-through on the importer's dollar invoicing share. The posterior median indicates that the importer's dollar invoicing share explains 15% of the cross-dyad variance in dollar pass-through, with 95% equal-tailed posterior credible interval [7.1%, 24.6%]. Thus, knowing the importer's country-level dollar invoicing share substantially improves the ability to explain cross-dyad heterogeneity in price pass-through, as predicted by DCP.

5.3 Results: trade elasticity

Our explorations of the elasticity of trade quantities suggest that their cross-dyad heterogeneity with respect to the dollar exchange rate is also related to the dollar invoicing share. Appendix B.3.1 provides the details. In a nutshell, our empirical specification again follows Section 4 with $Y_{ij,t} = \Delta y_{ij,t}$ in Eq. (8) with one lag of bilateral and dollar exchange rates, as well as the contemporaneous value and lag of importer log real GDP growth as controls. We find that the conditional density of the dollar trade elasticity (expected to be a negative number, as estimated in Section 4) shifts leftward when the importer's country-level dollar invoicing share increases. That is, the higher the dollar invoicing share, the larger is the average dollar trade elasticity in absolute value. However, our estimates on the trade elasticity are generally associated with higher posterior uncertainty than those for the price pass-through.

6 Model

The strong empirical support for DCP implies that U.S. monetary policy could play a special role. In this section, we pursue this idea theoretically. We simulate a DSGE model with three large countries/regions, U, G, and R. International prices are sticky in a dominant currency, namely the currency of country U. We consider three countries as opposed to two countries (the typical set-up in PCP and LCP environments) because of the non-standard predictions that can arise under DCP for trade and other spillovers between non-dominant currency countries.

Our goal is to examine the differential global implications of monetary policy (MP) shocks originating in the dominant currency country (U) relative to shocks originating in the non-dominant currency countries (G or R). We demonstrate that, when monetary policy is set using a Taylor rule: (i) MP shocks in U have *strong* spillovers to MP in G and R. On the other hand, MP shocks in G/R have *weak* spillovers to MP in U; (ii) A tighter MP in U and the accompanying appreciation of the dollar *reduce* rest-of-world and global trade, while this is not true for monetary tightenings originating in other countries.

6.1 Model set-up

Our model is closely related to that of Casas et al. (2016), with the main difference being that here the countries are large and shocks in one country transmit to others. This contrasts with the small open economy case studied in Casas et al. (2016), where variables in the rest of the world are exogenous by construction. We describe the model details and calibration in Appendix B.4, and only discuss the impulse responses in this section.

Briefly, the model has three regions that are symmetric in all respects, except for dollar pricing of internationally traded goods and the denomination of all internationally traded bonds in dollars. The latter two assumptions capture important features of the dollar's dominance both in international trade and finance, though admittedly in extremity, but help make our points most starkly.[21] Each country is made up of three types of agents: households, producers of non-traded goods, and producers of intermediate goods. Households save and provide labor monopolistically, and wage setting is subject to a Calvo friction. Producers of intermediate goods combine labor and intermediate inputs to produce a unique variety of good that is sold to the producer of

[21]We simulate the model also for the case when there is a full set of Arrow-Debreu securities traded. The impulse responses, qualitatively and quantitatively, are very close. This is intuitive because under perfect foresight, the noncontingent bond is sufficient to complete the market, i.e., the equilibrium conditions of the cases with complete markets and incomplete markets with a bond are the same. When an unanticipated shock hits, only the initial period's equilibrium conditions differ across the two cases.

non-traded goods. Price setting is also subject to a Calvo friction, and we assume that all goods sold domestically are priced in domestic currency and those sold internationally are priced in dollars as in DCP. The non-traded sector is perfectly competitive. This sector aggregates all domestic and imported varieties of intermediate goods to produce a good that is consumed as a final good and used as an intermediate input for production. The aggregator is defined by a Kimball (1995) homothetic demand aggregator which generates a role for strategic complementarity in pricing among producers of intermediate goods. Monetary policy follows a Taylor rule with inertia.

6.2 Monetary policy shock in U

We first consider a positive 25 basis point shock to the nominal interest rate in U. The impulse responses to this monetary tightening in U are plotted in Fig. 11. The outcomes in G and R are the same for all variables, including their exchange rates, both of which depreciate by 0.65% relative to the dollar on impact.

Impact on U. The rise in interest rates in U leads to a decline in output (-0.6%, Fig. 11(e)) and consumption (-0.36%, Fig. 11(f)), and a fall in inflation (-0.02%, Fig. 11(d)). The decline in inflation is, however, negligible (in contrast to PCP) because dollar pricing generates a low pass-through of the dollar appreciation into the price of imported goods, as seen in Fig. 11(k). On the other hand, the pass-through into export prices (in the destination currency) is high, as depicted in Fig. 11(l), which in turn generates a significant decline in exports (Fig. 11(m)). Imports decline because of the decline in overall demand given MP tightening. The trade balance to GDP, plotted in Fig. 11(i), deteriorates mildly. The terms of trade are largely unchanged.

Impact on G/R. The monetary tightening in U has a larger effect on inflation on impact in G/R (0.2%, Fig. 11(d)) than in U because the depreciation has high pass-through into import prices of the former countries. This in turn generates an endogenous increase in interest rates (0.15%, Fig. 11(b)) in G/R via the Taylor rule, leading to a mild contraction in output (-0.03%, Fig. 11(e)) and consumption (-0.13%, Fig. 11(f)) in G/R. Despite the depreciation of the G/R exchange rates relative to the dollar, their exports to U decline (-0.4%, Fig. 11(n)) because dollar prices to U change by little so there is no significant positive expenditure switching effect, and the decline in overall demand in U generates a decline in exports to U. Also, because of dollar pricing, there is a sharp decline in exports from G to R (-0.85%, Fig. 11(n)) and vice versa. This is because the depreciation of these countries' currencies relative to the dollar makes all imports more expensive, leading to a switch in expenditures away from imported goods. This is then further accentuated by the (mild) negative impact on consumption from the rise in interest rates in response to the inflationary effect.

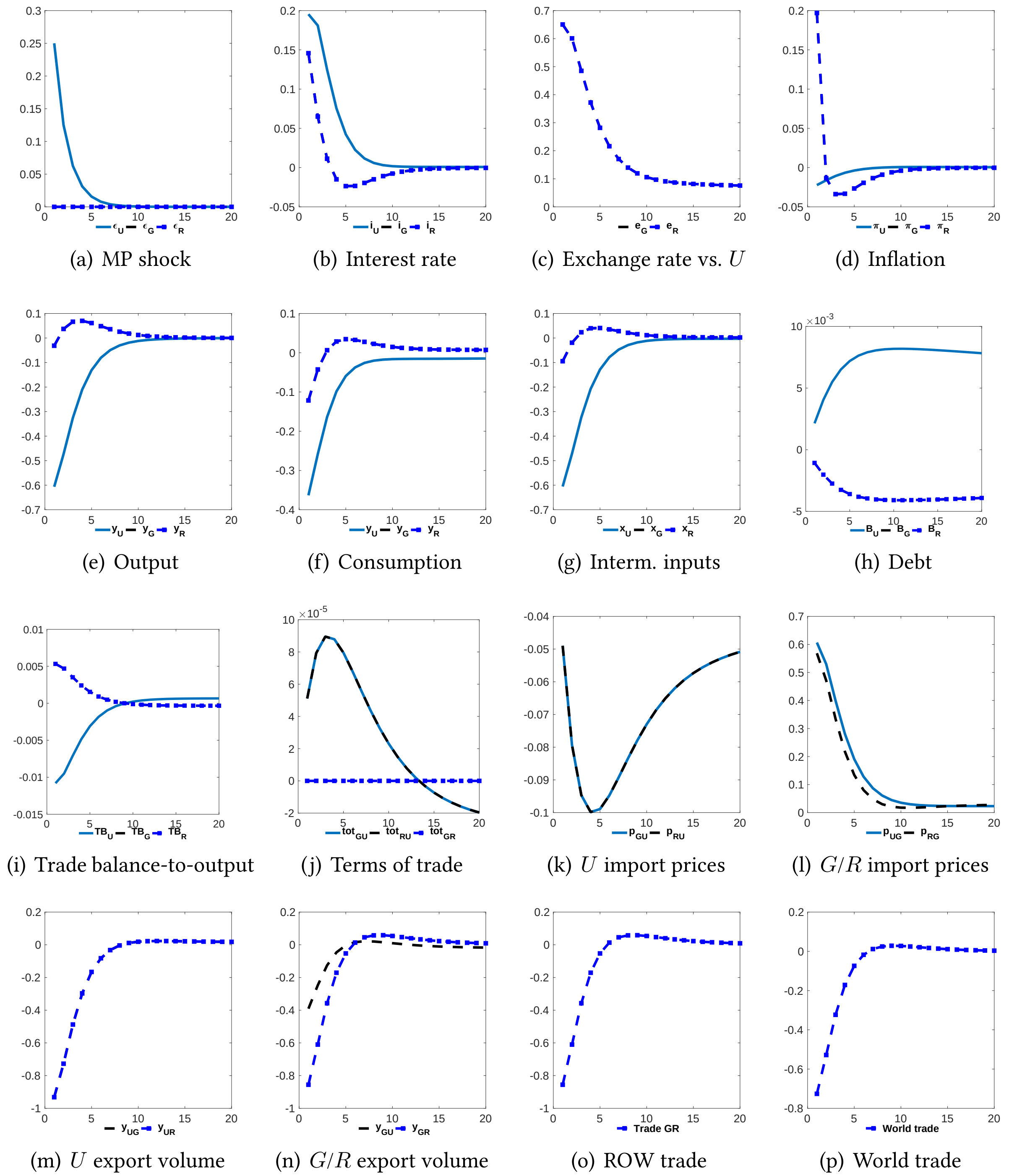
Impact on global trade. As follows from the previous discussion, a monetary tightening in U and the accompanying uniform appreciation of the dollar relative to other countries generate a decline in rest-of-world trade (-0.83%, Fig. 11(o)), defined as the sum of quantities traded between G and R. It also causes a decline in global trade (-0.73%, 11(p)), defined as the sum of export quantities from all countries.

6.3 Monetary policy shock in G/R

We next consider a 25 basis point monetary tightening in a non-dominant currency country. Without loss of generality, we set this to be G. As depicted in Fig. 12(c), G's currency appreciates uniformly relative to U and R on impact, and by a magnitude similar to that in Fig. 11(c). This is because, despite the endogenous change in interest rates in each country (Fig. 12(b) differs from Fig. 11(b)), the change in the interest rate differential between countries is quite similar, which is what matters for the exchange rate change.

Impact on G. The transmission of the shock to interest rates in G (Fig. 12(b)) is partly muted because the decline in inflation is endogenously contained through the Taylor rule. The negative impact on inflation of -0.2% (Fig. 12(d)) contrasts with the much smaller effect of a MP shock in U on U's inflation. This differential response arises from the strong pass-through of the appreciation of G's currency into its import prices. The rise in interest rates in G leads to a decline in output (-0.6%, Fig. 12(e)) and consumption (-0.27%, Fig. 12(f)). While pass-through into import prices (in G's currency) is high, pass-through into export prices (in destination currency) is low. Consequently, there is only a small negative impact on exports from G, in contrast to the large negative impact of a MP tightening in U on U's exports. While exports are not responsive, there is a significant increase in imports into G from U and R through the expenditure switching channel following the depreciation of their currencies relative to G's. The terms of trade are stable, as in the case of the MP shock in U.

Impact on global trade. The monetary tightening in G is associated with an expansion in global trade and almost no effect on rest-of-world trade (gross trade between U and R). Exports from U and R to G increase significantly, while exports out of G decline only marginally. Consequently, global trade as a whole expands. Even though the exchange rate between U and R



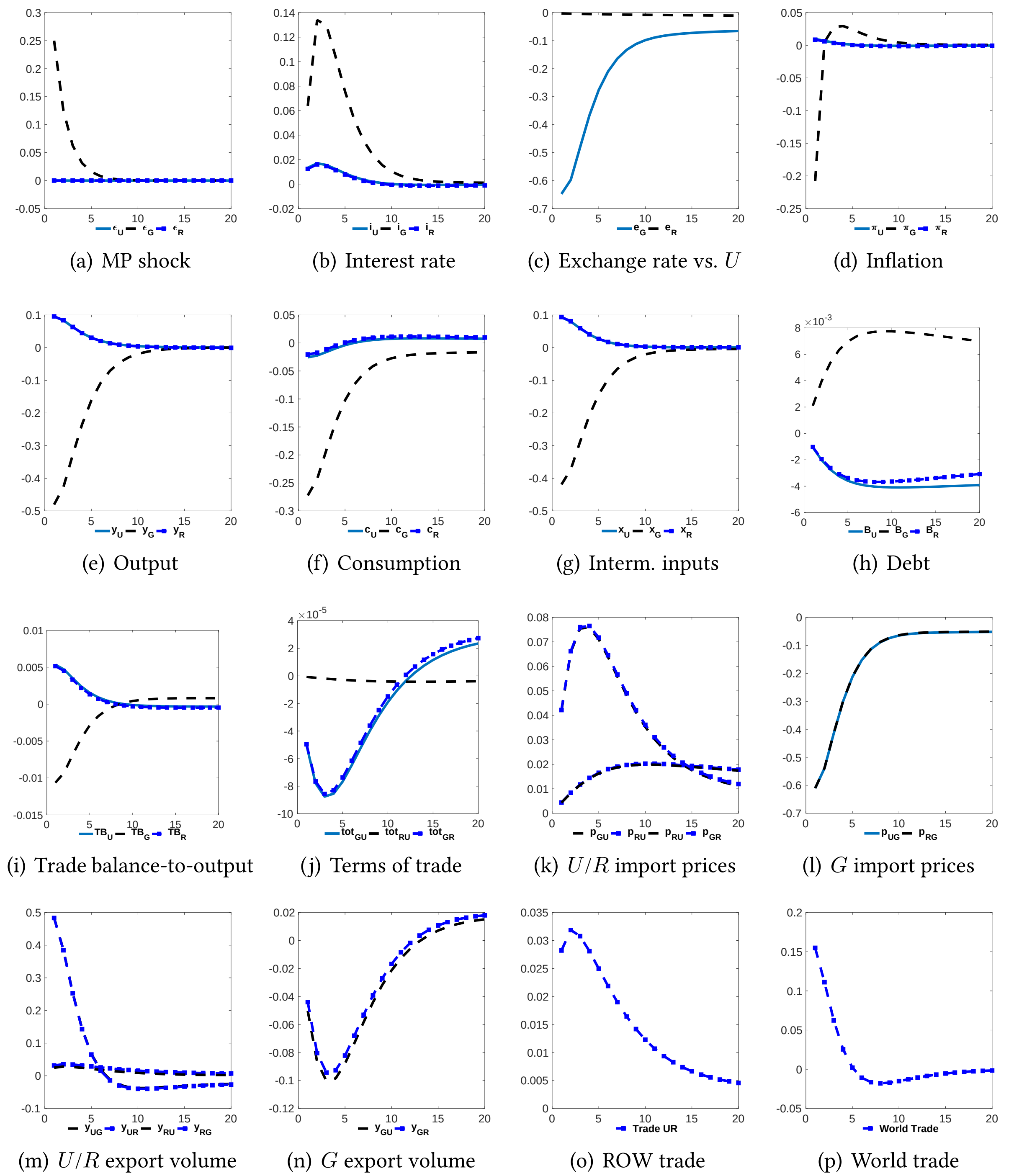
(a) MP shock (b) Interest rate (c) Exchange rate vs. U (d) Inflation

(e) Output (f) Consumption (g) Interm. inputs (h) Debt

(i) Trade balance-to-output (j) Terms of trade (k) U import prices (l) G/R import prices

(m) U export volume (n) G/R export volume (o) ROW trade (p) World trade

Figure 11: Impulse responses to a 25 basis point monetary tightening in U. Rest-of-world trade is defined as the sum of quantities traded between G and R. World trade is defined as the sum of export quantities from all countries.



(a) MP shock (b) Interest rate (c) Exchange rate vs. U (d) Inflation

(e) Output (f) Consumption (g) Interm. inputs (h) Debt

(i) Trade balance-to-output (j) Terms of trade (k) U/R import prices (l) G import prices

(m) U/R export volume (n) G export volume (o) ROW trade (p) World trade

Figure 12: Impulse responses to a 25 basis point monetary tightening in G. Rest-of-world trade is defined as the sum of quantities traded between U and R. World trade is defined as the sum of export quantities from all countries.

remains unchanged, the expansionary effect of the depreciation of their currencies relative to G and the commensurate increase in their outputs are associated with higher purchases of intermediate inputs from each other. This in turn results in a small increase in gross trade between the two countries.

7 Conclusion

Using newly constructed trade price and volume indices for over 2,500 country pairs, we document that the relevant predictor for bilateral trade prices and volumes is not the bilateral exchange rate but the dollar exchange rate, even when the U.S. is on neither side of the trade transaction. A 1% U.S. dollar appreciation against all other currencies in the world predicts a 0.6% decline within a year in the volume of total trade between countries in the rest of the world, controlling for the global business cycle. We also demonstrate that the impact of the dollar exchange rate is increasing in the fraction of a country's trade that is invoiced in dollars. We estimate that the importing country's share of imports invoiced in dollars explains 15% of the variance of dollar pass-through across country pairs. These findings strongly support DCP as the empirically relevant framework for understanding the international transmission of shocks and for policy analysis. As an illustration, our calibrated DSGE model demonstrates that DCP predicts stark differences in monetary policy spillovers originating from shocks in dominant vs. non-dominant currency countries.

On a methodological note, our Bayesian analysis demonstrates the ease with which rich hierarchical econometric models can be estimated with the user-friendly open source software Stan. We expect that semiparametric hierarchical panel data analysis will prove useful also in other empirical settings where understanding cross-sectional heterogeneity is of primary importance.

A Appendix

A.1 Data

Here we provide further details on the Comtrade, WDI, and FRED data.

Comtrade country summary statistics. Table 8 lists summary statistics on the number of observations for the 55 countries in our merged Comtrade/WDI dataset. The table also lists the share of imports invoiced in U.S. dollars and euros for the 39 countries for which we observe these measures (cf. Gopinath, 2015).

World Development Indicator data. The exchange rate is the World Bank's "alternative conversion factor" series (PA.NUS.ATLS), which corrects for redenominations and currency substitution, and is measured as an annual average of daily rates. Producer prices are given by the wholesale price index (FP.WPI.TOTL). Real GDP is measured at market prices in constant U.S. dollars (NY.GDP.MKTP.KD). The GDP deflator is given by the ratio of nominal GDP (NY.GDP.MKTP.CD) and real GDP. Consumer prices are constructed from CPI inflation rates (FP.CPI.TOTL.ZG), or if inflation is not available, CPI levels (FP.CPI.TOTL). We use data for 1989–2015 only.

FRED data. We obtain the WTI oil price (POILWTIUSDA), VIX (VIXCLS), and 1-year Treasury bill rate (DTB1YR) from the St. Louis Fed's FRED database. Annual series are averages of daily indices.

A.2 Bayesian analysis: priors and model selection

A.2.1 Hyper-priors

Here we describe the remaining parts of the prior not specified in the main text. We incorporate time fixed effects δ_t by adding $T-1$ dummies in the covariate vector X_t, so the parameter vector θ includes these parameters. We impose the following priors, all mutually independent:

$$\alpha \sim \text{Cauchy}(0,5), \quad \theta_j \sim \text{Cauchy}(0,5),$$

$$\sigma \sim \text{HalfCauchy}(0,1), \quad \tau \sim \text{HalfCauchy}(0,1).$$

Cauchy$(0,a)$ is the centered Cauchy distribution with interquartile range $2a$. HalfCauchy$(0,a)$ is the restriction of the Cauchy$(0,a)$ distribution to the positive real line. Since the units of our

Country summary statistics

Country	Adv	As exporter #dyads	As exporter avg T	As importer #dyads	As importer avg T	InvS$^{\$}$	InvS$^{€}$
Africa							
Algeria		20	12.9	46	20.9		0.49
Egypt		53	20.2	50	18.0		
South Africa		51	14.8	53	14.7		
Americas							
Argentina		54	21.0	50	20.6	0.88	0.08
Brazil		54	21.7	50	23.2	0.84	0.11
Canada	✓	54	22.0	53	24.2	0.75	0.05
Chile		52	20.2	48	17.7		
Colombia		52	17.9	49	15.6	0.99	0.00
Mexico		54	21.7	51	23.0		
United States	✓	54	22.0	53	22.8	0.93	0.02
Venezuela		8	17.6	46	17.0		
Asia							
China		54	21.9	53	21.7		
Hong Kong	✓	53	22.1	51	20.7		
India		54	21.9	53	24.0	0.86	0.10
Indonesia		53	21.6	51	21.8	0.81	0.04
Israel	✓	49	22.1	50	15.0	0.73	0.21
Japan	✓	54	22.1	52	25.4	0.71	0.03
Kazakhstan		32	15.2	52	14.6		
Malaysia		53	22.0	50	23.8		
Philippines		54	21.6	47	18.0		
Saudi Arabia		50	19.7	50	15.3		
Singapore	✓	54	22.0	50	23.6		
South Korea	✓	54	22.0	51	23.7	0.81	0.05
Thailand		54	21.8	51	24.7	0.79	0.04
Turkey		54	22.0	52	24.0	0.59	0.31
Vietnam		50	19.6	46	12.1		

(continued on next page)

Country summary statistics (continued)

Country	Adv	As exporter		As importer			
		#dyads	avg T	#dyads	avg T	InvS$^{\$}$	InvS$^{€}$
Europe							
Austria	✓	54	22.2	52	20.7	0.06	0.70
Belgium	✓	53	15.8	53	15.9	0.14	0.82
Czech Republic	✓	53	20.2	53	21.2	0.19	0.68
Denmark	✓	54	22.0	52	24.2	0.25	0.32
Estonia	✓	46	17.0	52	18.0	0.34	0.53
Finland	✓	54	21.9	52	24.9	0.42	0.38
France	✓	54	22.2	53	20.7	0.21	0.75
Germany	✓	54	21.4	53	23.3	0.23	0.75
Greece	✓	54	21.4	51	22.0	0.40	0.58
Hungary		54	22.0	52	21.5	0.27	0.57
Ireland	✓	54	21.9	52	21.7	0.23	0.47
Italy	✓	54	22.2	52	20.7	0.29	0.67
Lithuania	✓	51	16.8	48	19.0	0.51	0.39
Luxembourg	✓	49	15.6	51	13.6	0.16	0.78
Netherlands	✓	54	22.2	53	22.2	0.37	0.46
Norway	✓	54	22.0	51	21.6	0.21	0.29
Poland		54	21.8	52	20.2	0.30	0.58
Portugal	✓	54	21.8	52	25.0	0.22	0.76
Romania		53	21.1	50	19.7	0.31	0.67
Russia		53	21.0	52	17.6		
Slovak Republic	✓	50	18.9	51	20.0	0.12	0.79
Slovenia	✓	54	19.6	52	20.0	0.20	0.75
Spain	✓	54	22.0	54	24.8	0.35	0.58
Sweden	✓	54	22.0	54	21.9	0.25	0.36
Switzerland	✓	54	22.1	54	25.1	0.13	0.53
Ukraine		51	18.8	52	17.2	0.75	0.16
United Kingdom	✓	54	22.2	54	21.6	0.47	0.15
Oceania							
Australia	✓	54	21.8	51	25.4	0.53	0.08
New Zealand	✓	53	20.7	50	23.5		

Table 8: Summary statistics for countries in the merged Comtrade/WDI sample. Adv: advanced economy (IMF WEO). #dyads: number of non-missing dyads that the country appears in. avg T: average number of years per dyad that the country appears in; a dyad-year observation is counted if at least one UVI or volume observation is reported by the importer, and exchange rate data exists for both countries. InvS: share of imports invoiced in USD/euro.

outcome variables $Y_{ij,t}$ are log points, the above priors are highly diffuse. As for the MGLR prior, we assume[22]

$$\omega_k \sim \text{HalfCauchy}(0,2), \qquad \begin{pmatrix} \mu_{0,k} \\ \mu_{1,k} \end{pmatrix} \mid \omega_k \sim N\left(0, \begin{pmatrix} \omega_k^2 & 0 \\ 0 & \omega_k^2 \end{pmatrix}\right), \qquad k = 1, \dots, K,$$

$$\zeta_k(\cdot) \sim GP(0, C(\cdot; A_k)), \quad A_k \sim \text{Exponential}(1), \quad k = 1, \dots, K-1,$$

independently across k. Here $GP(0, C(\cdot; A))$ denotes a Gaussian process with Gaussian radial covariance kernel

$$C(s_1, s_2; A) = \exp\{-A(s_1 - s_2)^2\} + 0.0001 \times \mathbb{1}(s_1 = s_2), \quad s_1, s_2 \in [0,1].$$

The second term on the right-hand side above helps avoid numerical issues in the warm-up phase of the MCMC algorithm, but it is small enough to negligibly affect the final output (the dollar invoicing share S_j is measured as a fraction between 0 and 1).

A.2.2 Bayesian leave-one-out cross-validation

The Bayesian Leave-One-Out (LOO) cross-validation criterion of Gelfand et al. (1992) is given by the cross-sectional sum of leave-one-out predictive densities

$$\begin{aligned} LOO &= \sum_{ij} \log f(Y_{ij} \mid R_{ij}, X_{ij}, Y_{-(ij)}, R_{-(ij)}, X_{-(ij)}) \\ &= \sum_{ij} \log \int f(Y_{ij} \mid R_{ij}, X_{ij}, \vartheta)\, f(\vartheta \mid R_{ij}, X_{ij}, Y_{-(ij)}, R_{-(ij)}, X_{-(ij)})\, d\vartheta. \end{aligned}$$

Here ϑ collects all model parameters. $Y_{ij} = (Y_{ij,1}, \dots, Y_{ij,T})$ collects all observed outcomes for dyad (i,j) across time, and similarly for the covariates R_{ij} and X_{ij}.[23] The notation $Y_{-(ij)}$ means all observed outcomes for dyads other than (i,j), and similarly for $R_{-(ij)}$ and $X_{-(ij)}$. The LOO criterion is large when the model yields good (leave-one-out) out-of-sample fit, given knowledge of the covariates. This is similar in spirit to the well-known non-Bayesian leave-one-out cross-validation criterion. We use a Pareto-smoothed importance sampling estimate of LOO, as developed by Vehtari et al. (2017) and implemented in Stan.

[22]Because the mixture component labels are not identified, we additionally impose the normalization $\mu_{0,1} < \mu_{0,2} < \cdots < \mu_{0,K}$. Stan accomplishes this by reparametrizing the vector $(\mu_{0,1}, \dots, \mu_{0,K})'$ into an unconstrained parameter, while adjusting for the Jacobian of the transformation in the posterior density.

[23]Since we have an unbalanced panel, the dimension of Y_{ij}, R_{ij}, X_{ij} actually varies across dyads.

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B Online Appendix (NOT FOR PUBLICATION)

B.1 Data: Comparison of Comtrade and BLS price series for the U.S.

In this section we compare our unit value indices to survey price indices from the U.S. Bureau of Labor Statistics. The BLS provides U.S. import price indices by locality of origin for Canada, E.U., France, Germany, U.K, Latin America, Mexico, Pacific Rim, China, Japan, ASEAN, Asia Near East, and Asian Newly Industrialized countries. As these price indices are constructed from surveys, their comparison with our unit value based indices can help gauge the effectiveness of our techniques to deal with the unit value bias and other potential mismeasurement inherent in customs data.

To arrive at comparable series, in this subsection we follow BLS in using *Laspeyres* indices of *total* (commodities and non-commodities) goods prices from our Comtrade data set. For regions with multiple countries, we aggregate country level growth rates using Comtrade import values with a two year lag. Still, the series are not fully comparable because BLS' preferred price basis is f.o.b. (free on board) while import values recorded at customs are c.i.f. (cost, insurance and freight), and not all countries included in BLS regions are in our database.

Our indices constructed from Comtrade unit values track the BLS import price indices fairly well, as shown in Figs. 13 and 14. These figures compare the linearly detrended logged indices, since our regressions use log growth rates and absorb any disparity in average growth rates in the intercept. The growth rates of our indices for Canada, Japan, Mexico, and the aggregated Latin America and Asia Near East match those of BLS remarkably well. The comparison with some Asian countries suggests that a unit value bias may still be present, causing the unit value series to be somewhat more volatile than the BLS price series. Nevertheless, for every country group and individual country except Germany, the correlation coefficient between the Comtrade and BLS growth rates is high. Finally, the match for European countries seems acceptable, with the year 2008 being an exception. A closer inspection of the case of Germany reveals that a couple of products (transport vehicles) with large import shares experienced substantial unit value decreases that year according to Comtrade, leading our indices to decline while the BLS index shows an increase.

Comtrade and BLS import price indices for U.S.: country groups

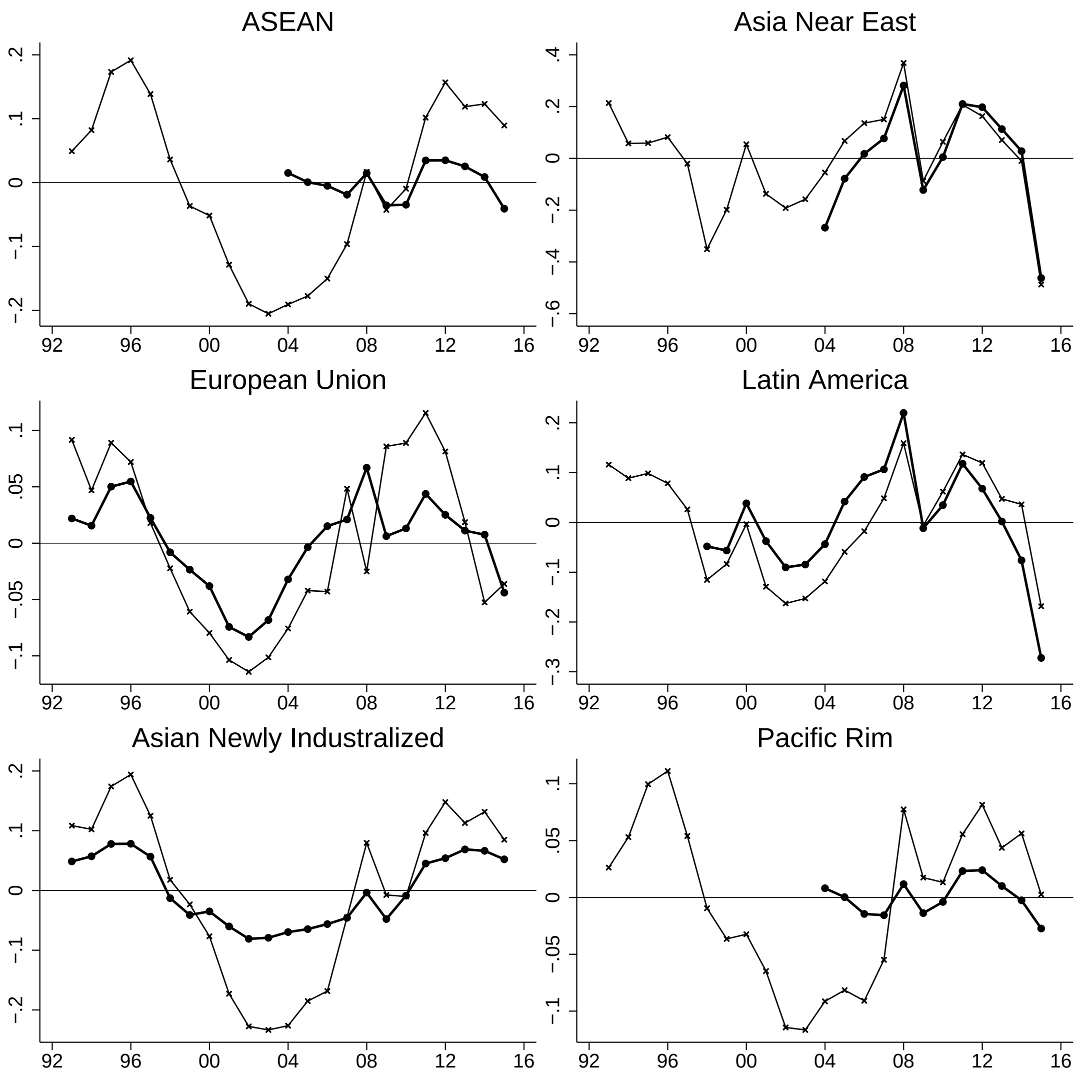

Figure 13: Comparison of BLS Locality of Origin import price indices (thick lines, circles) with our constructed Comtrade analogues (thin lines, crosses). Plotted indices are logged and linearly detrended. The Comtrade sample does not cover all countries in the BLS country groups, cf. Table 9.

Comtrade and BLS import price indices for U.S.: individual countries

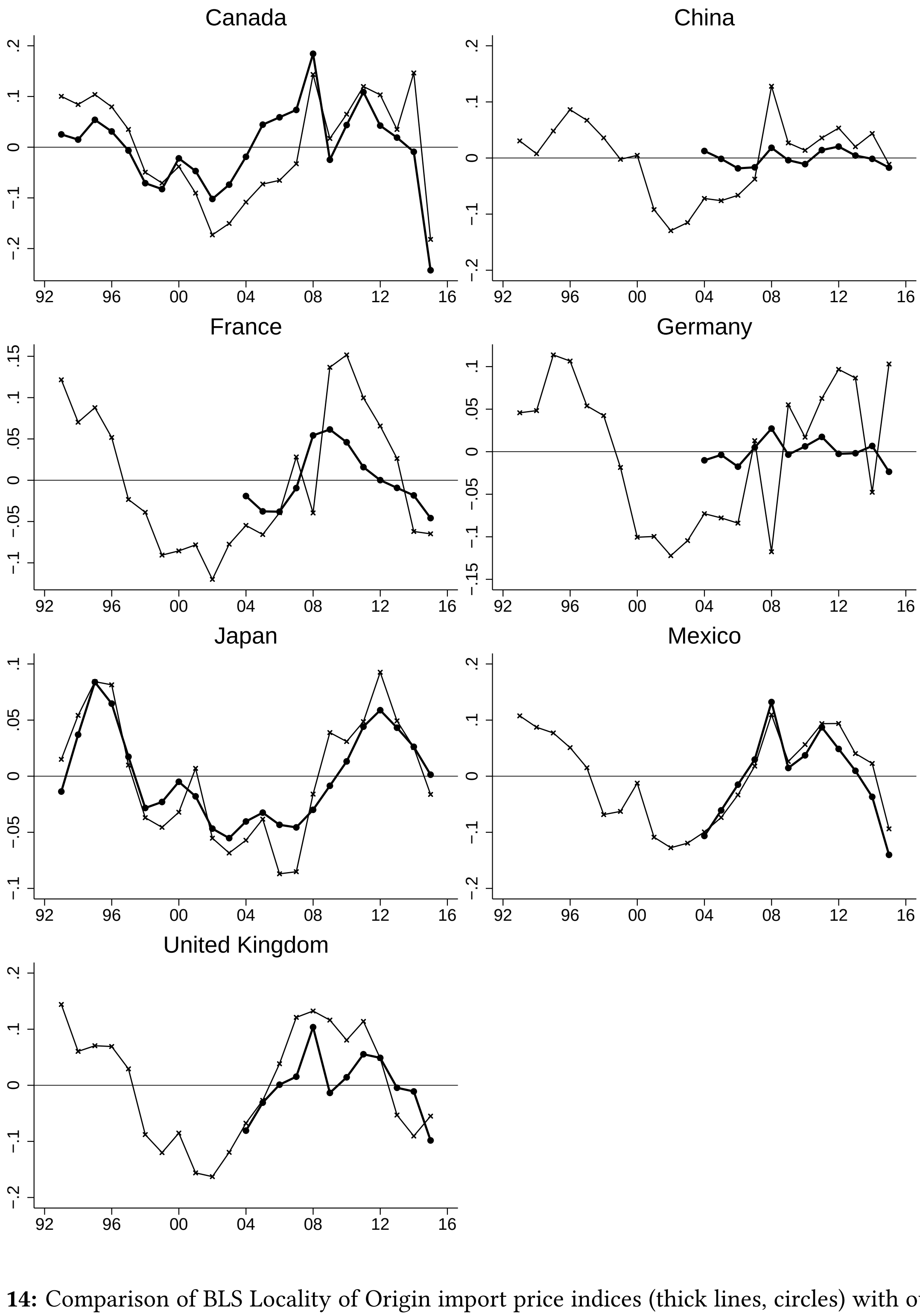

Figure 14: Comparison of BLS Locality of Origin import price indices (thick lines, circles) with our constructed Comtrade analogues (thin lines, crosses). Plotted indices are logged and linearly detrended.

BLS COUNTRY GROUPS

BLS group	Country ISO codes
ASEAN	BRN* IDN KHM* LAO* MMR* MYS PHL SGP THA VNM*
Asia Near East	ARE* BHR* IRN* IRQ* ISR JOR* KWT* LBN* OMN* QAT* SAU SYR* YEM*
European Union	AUT BEL BGR* CYP* CZE DEU DNK ESP EST FIN FRA GBR GRC HRV* HUN IRL ITA LTU LUX LVA* MLT* NLD POL PRT ROU SVK SVN SWE
Latin America	ARG BRA CHL COL MEX VEN (plus other unspecified Central American, South American, and Caribbean countries*)
Asian New. Ind.	HKG KOR SGP TWN
Pacific Rim	AUS BRN* CHN HKG IDN JPN KOR MAC* MYS NZL PHL PNG* SGP TWN

Table 9: Definition of BLS country groups in Fig. 13. Countries marked with an asterisk (*) are not available in the Comtrade sample.

B.2 Panel regressions: Supplementary results

This section provides supplementary panel regression results, including robustness checks.

B.2.1 Spillovers from U.S. dollar to foreign inflation

Our results imply that fluctuations in the strength of the dollar, for example those caused by U.S. monetary policy actions, have spillover effects on foreign inflation. We have shown that the dollar exchange rate passes strongly through to bilateral import prices measured in the importer's currency, especially for countries whose imports are heavily invoiced in dollars. Given a non-negligible import content in consumption, this implies that dollar movements will directly affect foreign consumer price index (CPI) inflation, as discussed by Gopinath (2015). If foreign firms behave in a monopolistically competitive way, foreign producer prices will react to changes in foreign import prices, although perhaps with a lag. Hence, the direct effect of dollar movements on foreign CPI inflation may be amplified by endogenous producer responses.

We now provide direct country-level regression evidence on the effects of the U.S. dollar exchange rate on foreign consumer and producer prices. Gopinath (2015) computes back-of-the-envelope estimates of these spillovers based on estimated country-level import price pass-through and the import content of consumption. We instead directly regress countries' CPI or PPI on the dollar exchange rate. Additionally, we investigate the interaction of the dollar exchange rate and the dollar import invoicing share.

Specifically, we consider the country-level panel regression

$$\Delta cpi_{j,t} = \lambda_j + \delta_t + \sum_{k=0}^{2} \beta_k \Delta e_{\$j,t-k} + \sum_{k=0}^{2} \psi_k \Delta e_{\$j,t-k} \times S_j + \varepsilon_{j,t}, \tag{10}$$

where $\Delta cpi_{j,t}$ is the change in the log CPI in the currency of country j, and λ_j and δ_t are country and year fixed effects, respectively. We also consider specifications with $\Delta ppi_{j,t}$ on the left-hand side, as well as specifications restricting $\psi_k = 0$ for all k. We focus attention on the post-2002 (post-euro) sample, since full-sample regression results are unduly influenced by a handful of countries' high-inflation/high-depreciation episodes in the 1990s.[24]

Table 10 displays the contemporaneous dollar pass-through into CPI and PPI. The first two columns shows results for CPI pass-through, and the second two show those for PPI pass-through. Columns (1) and (3) do not interact exchange rate changes with the dollar invoicing share, while

[24]The results are very similar if we use the full 1992–2015 sample but drop country-year observations for which the arithmetic CPI inflation rate exceeds 30% annually (0.26 log inflation rate).

Dollar pass-through into CPI and PPI, 2002–2015

VARIABLES	(1) $\Delta cpi_{j,t}$	(2) $\Delta cpi_{j,t}$	(3) $\Delta ppi_{j,t}$	(4) $\Delta ppi_{j,t}$
$\Delta e_{\$j,t}$	0.106***	0.0221	0.284***	0.182***
	[0.04, 0.18]	[-0.05, 0.09]	[0.14, 0.43]	[0.05, 0.32]
$\Delta e_{\$j,t} \times S_j$		0.181**		0.237*
		[0.04, 0.33]		[-0.03, 0.51]
ΔER lags	2	2	2	2
Time FE	yes	yes	yes	yes
R-squared	0.283	0.453	0.532	0.675
Observations	766	544	697	525
Countries	55	39	52	38

Table 10: The first (resp., last) two columns use CPI (resp., PPI) growth as dependent variable. 95% confidence intervals clustered by country and corrected for small number of clusters using "LZ2-BM" method of Imbens and Kolesár (2016). *** p<0.01, ** p<0.05, * p<0.1 (only 95% interval shown).

columns (2) and (4) do. The table displays 95% confidence intervals rather than standard errors because the small number of countries (clusters) necessitates the use of small-sample corrections (Imbens and Kolesár, 2016). The average pass-through of the dollar into CPI (resp., PPI) is 11% (resp., 28%) within the year. The dollar pass-through is larger for countries that have a substantial fraction of imports invoiced in dollars. The contemporaneous interaction term is statistically significant at the 10% level for both the CPI and PPI specifications, and also at the 5% level for the CPI specification. Hence, it appears that countries which invoice more in dollars experience higher dollar pass-through into consumer and producer prices. We caution, though, that the *magnitude* of the pass-through is imprecisely estimated when controlling for country and time fixed effects.

B.2.2 Regression details and robustness checks

Post-2002 results. Exchange rate pass-through into prices has been stable over our sample period, while trade elasticities may have become larger in absolute value in the latter part of the sample. We compute results for the subsample 2002–2015, roughly corresponding to the second half of our data set, and also corresponding to the sample used for the euro regressions in Section 4.5. Figs. 15 and 16 show price and volume impulse responses for the 2002–2015 subsample that correspond to the full-sample results in Figs. 1 and 3 in Section 4. The price pass-through impulse responses of bilateral and dollar exchange rates are similar to the full-sample results. However, the post-2002 USD cumulative trade elasticity (unweighted) is substantially negative at lags of 1 and 2 years, whereas the level effect is close to zero at lags 1 and 2 on the full sample.

Pre-2007 results. Our headline results are not driven by the global financial crisis starting in 2008. Figs. 17 and 18 show the average exchange rate pass-through and trade elasticity computed on the 1992–2007 sample. The results are almost identical to our baseline Figs. 1 and 3.

Euro regressions. Tables 11 and 12 display the results of the price pass-through and trade elasticity regressions in Section 4.5 involving the euro exchange rate. The regressions do not control for time fixed effects but do include the aggregate control variables listed in Section 4.5. Specifications (1) and (4) focus on the bilateral and euro exchange rates, specifications (2) and (5) add the dollar exchange rate, and specifications (3) and (6) include interactions with the dollar and euro import invoicing shares. $S_j^{€}$ is the importing country's share of imports invoiced in euros from Gopinath (2015). The interactions are statistically and economically significant and mostly have the expected signs in the price pass-through regressions: A higher share of euro (resp., dollar) invoicing implies a higher pass-through from the euro (resp., dollar) exchange rate.

Weighted average dollar invoicing share. Fig. 19 depicts the weighted average dollar import invoicing share $\sum_{j \neq \text{US}} w_j S_j$ used in Section 4.6, where the ex-U.S. non-commodity import value weights w_j have been computed for each year in our sample. Notice that the weighted average fluctuates tightly around a mean of 0.40.

Additional controls. Table 13 shows that our pass-through regressions results are qualitatively robust to adding importer PPI growth and importer real GDP growth as additional controls. We use two lags of the log changes of each of these indices. Although our baseline specification in Section 4.2 is common in the literature, the addition of importer PPI and GDP controls can be justified by models with strategic complementarity in pricing and country-specific demand shifts. While the overall level of both bilateral and USD pass-through is somewhat lower when the controls are added, our qualitative conclusions regarding the dominance of the USD exchange rate and the relationship with dollar invoicing are as pronounced in Table 13 as in Table 3.

Average price pass-through, 2002–2015

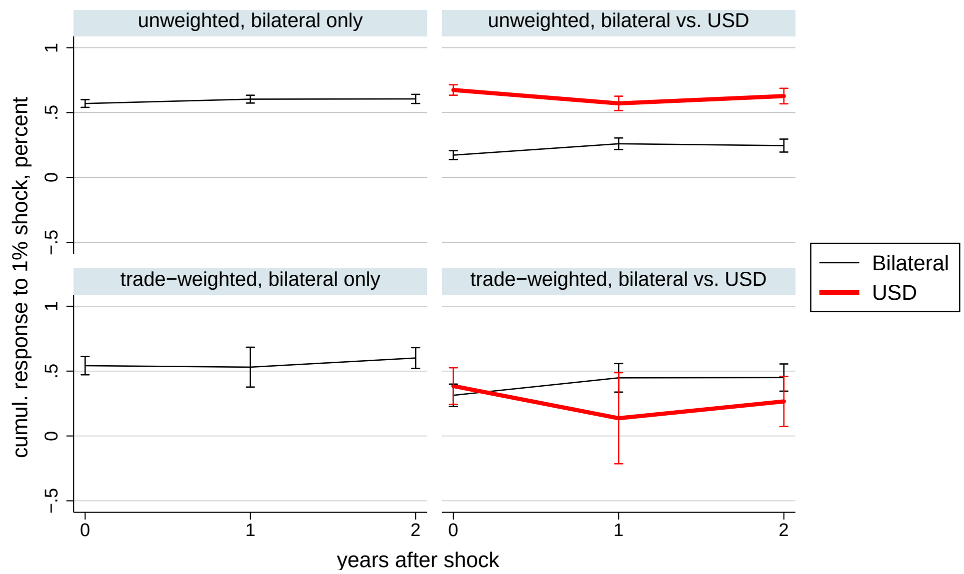

Figure 15: Fig. 1 computed on post-2002 data, but with same weights.

Average trade elasticity, 2002–2015

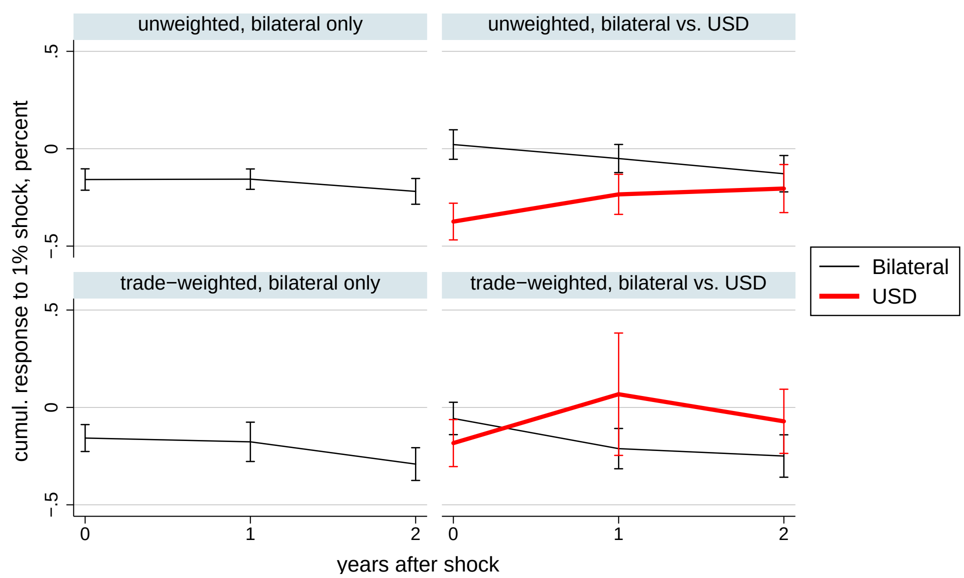

Figure 16: Fig. 3 computed on post-2002 data, but with same weights.

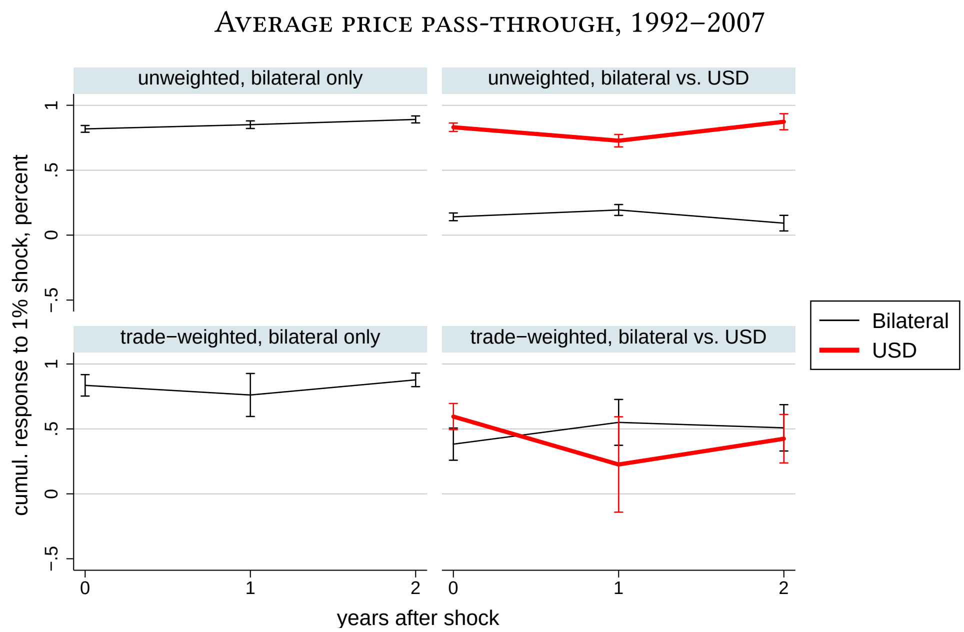

Figure 17: Fig. 1 computed on pre-2007 data, but with same weights.

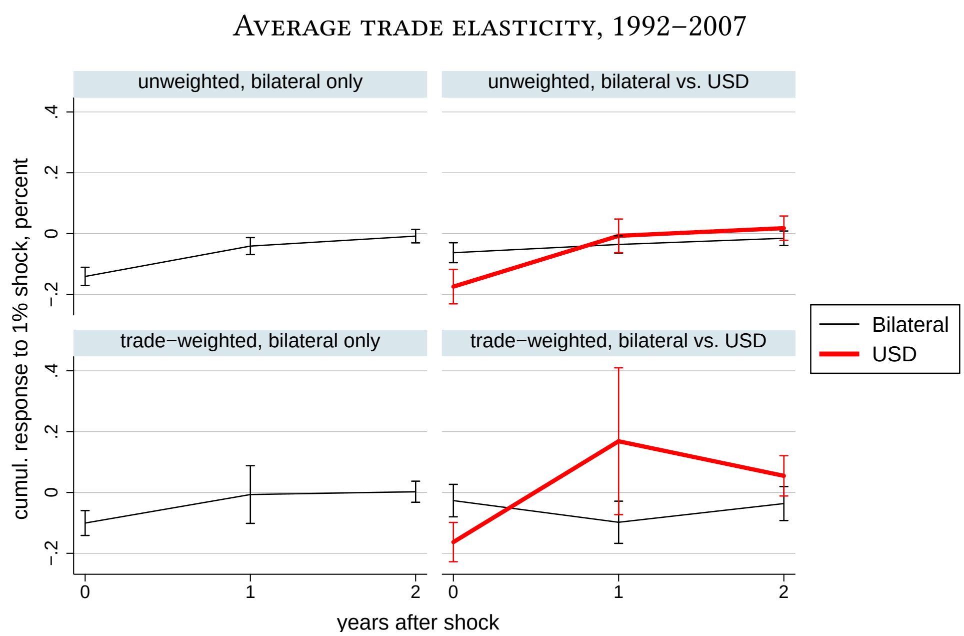

Figure 18: Fig. 3 computed on pre-2007 data, but with same weights.

Euro vs. dollar exchange rate pass-through into prices

	unweighted			trade-weighted		
	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$
$\Delta e_{ij,t}$	0.305***	0.181***	0.207***	0.438***	0.331***	0.551***
	(0.0159)	(0.0174)	(0.0695)	(0.0490)	(0.0567)	(0.156)
$\Delta e_{ij,t} \times (S_j + S_j^{€})$			-0.0357			-0.361**
			(0.0784)			(0.174)
$\Delta e_{\$j,t}$		0.754***	0.614***		0.561***	0.379***
		(0.0373)	(0.0405)		(0.0755)	(0.0672)
$\Delta e_{\$j,t} \times S_j$			0.510***			0.769***
			(0.0439)			(0.151)
$\Delta e_{€j,t}$	0.467***	-0.0800**	-0.347***	0.207***	-0.184***	-0.384***
	(0.0175)	(0.0332)	(0.0430)	(0.0612)	(0.0601)	(0.0726)
$\Delta e_{€j,t} \times S_j^{€}$			0.694***			0.709***
			(0.0821)			(0.122)
ΔER lags	2	2	2	2	2	2
Exp. PPI	yes	yes	yes	yes	yes	yes
Agg. controls	yes	yes	yes	yes	yes	yes
Time FE	no	no	no	no	no	no
R-squared	0.131	0.143	0.210	0.102	0.112	0.293
Observations	33,802	33,802	24,463	33,802	33,802	24,463
Dyads	2,647	2,647	1,900	2,647	2,647	1,900

Table 11: The first (resp., last) three columns use unweighted (resp. trade-weighted) regressions. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

Euro vs. dollar trade elasticity

VARIABLES	unweighted (1) $\Delta y_{ij,t}$	(2) $\Delta y_{ij,t}$	trade-weighted (3) $\Delta y_{ij,t}$	(4) $\Delta y_{ij,t}$
$\Delta e_{ij,t}$	-0.0631*	0.0229	-0.146***	-0.0560
	(0.0371)	(0.0386)	(0.0493)	(0.0429)
$\Delta e_{\$j,t}$		-0.695***		-0.573***
		(0.0806)		(0.124)
$\Delta e_{€j,t}$	-0.179***	0.320***	-0.00647	0.386***
	(0.0413)	(0.0759)	(0.0494)	(0.105)
ΔER lags	2	2	2	2
Imp. GDP	yes	yes	yes	yes
Agg. controls	yes	yes	yes	yes
Time FE	no	no	no	no
R-squared	0.068	0.071	0.197	0.203
Observations	37,437	37,437	37,437	37,437
Dyads	2,807	2,807	2,807	2,807

Table 12: The first (resp., last) two columns use unweighted (resp. trade-weighted) regressions. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

Weighted average dollar invoicing share over time

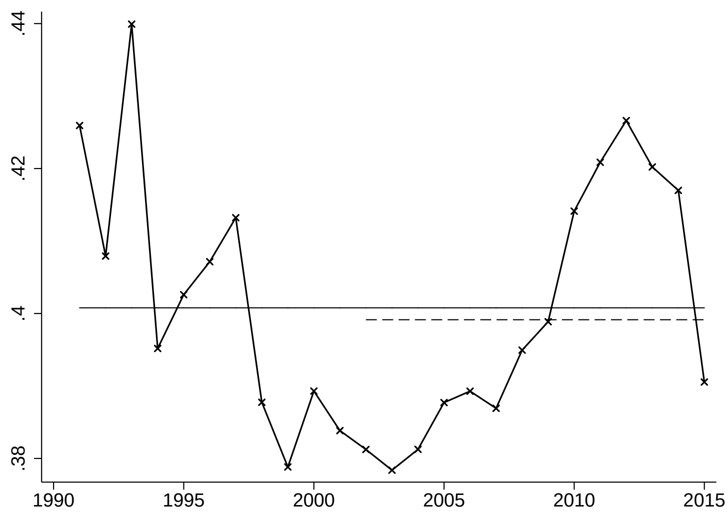

Figure 19: Weighted average dollar import invoicing share $\sum_{j\neq \text{US}} w_j S_j$, using import value weights w_j computed in different reference years (along horizontal axis). Horizontal lines show the mean on the 1992–2015 and 2002–2015 samples.

Exchange rate pass-through into prices: Additional controls

	unweighted			trade-weighted		
	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$	$\Delta p_{ij,t}$
$\Delta e_{ij,t}$	0.519***	0.163***	0.214***	0.550***	0.328***	0.456***
	(0.0117)	(0.0133)	(0.0177)	(0.0471)	(0.0480)	(0.0352)
$\Delta e_{ij,t} \times S_j$			-0.0869***			-0.272***
			(0.0252)			(0.0495)
$\Delta e_{\$j,t}$		0.706***	0.524***		0.464***	0.103
		(0.0183)	(0.0298)		(0.0347)	(0.0639)
$\Delta e_{\$j,t} \times S_j$			0.303***			0.643***
			(0.0360)			(0.0951)
ΔER lags	2	2	2	2	2	2
Exp. PPI	yes	yes	yes	yes	yes	yes
Imp. PPI	yes	yes	yes	yes	yes	yes
Imp. GDP	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes
R-squared	0.388	0.411	0.528	0.361	0.382	0.650
Observations	42,243	42,243	32,916	42,243	42,243	32,916
Dyads	2,502	2,502	1,853	2,502	2,502	1,853

Table 13: The first (resp., last) three columns use unweighted (resp. trade-weighted) regressions. S.e. clustered by dyad. *** p<0.01, ** p<0.05, * p<0.1.

B.3 Bayesian analysis: Supplementary results

This section provides supplementary results and implementation details for the Bayesian model.

B.3.1 Trade elasticity

Similar to the price pass-through results, we find that the cross-dyad heterogeneity of the elasticity of trade quantities with respect to the dollar exchange rate is related to the dollar invoicing share. However, the results in this subsection generally come attached with higher posterior uncertainty. Appendix B.3.2 provides additional results on parameters not highlighted below.

Our empirical specification again follows Section 4. We set $Y_{ij,t} = \Delta y_{ij,t}$ in Eq. (8). We control for one lag of bilateral and dollar exchange rates, as well as the contemporaneous value and lag of importer log real GDP growth. The sample of dyad-year observations is the same as for the price pass-through results.

We report results for $K = 4$ mixture components. The LOO model selection criterion strongly favors $K = 3, 4, 5$ against either $K \leq 2$ or $K = 6, 7, 8$. $K = 4$ has a slightly higher LOO score than $K = 3, 5$. However, we remark again that the results presented below are little changed across specifications with $K \geq 3$. We report results for $K = 8$ in Appendix B.3.3.

Fig. 20 shows that the conditional density of the dollar trade elasticity (expected to be a negative number, as also estimated in Section 4) shifts leftward when the importer's country-level dollar invoicing share increases. That is, the higher the dollar invoicing share, the larger in magnitude is the dollar trade elasticity, on average. Notice, however, that the credible bands are much wider here than for the price pass-through results. This is consistent with the larger standard errors on the interaction terms in the trade elasticity panel regressions in Section 4. Fig. 21 shows the conditional mean and standard deviation. While the posterior medians indicate that the conditional mean function is downward-sloping over most of the range of S_j, the function is estimated with substantial uncertainty.

Fig. 22 summarizes the posterior of the sample distribution of γ_{ij}. The median γ_{ij} is in line with the panel regression results in Section 4 (median median -0.11), but the heterogeneity is substantial (median IQR 0.09). Again we find a strong (here: negative, as expected) correlation between γ_{ij} and S_j (median correlation -0.41), after winsorizing γ_{ij} at 5% in each tail. Thus, trade elasticities with respect to the dollar are highly heterogeneous, but dyads with the largest-in-magnitude dollar elasticities tend to be the dyads with the highest importer dollar invoicing share. The 95% equal-tailed posterior credible interval for the R^2 in a cross-dyad regression of (winsorized) dollar elasticity on the importer's dollar invoicing share is [2.6%, 34.0%].

Density of dollar trade elasticity given dollar invoicing share

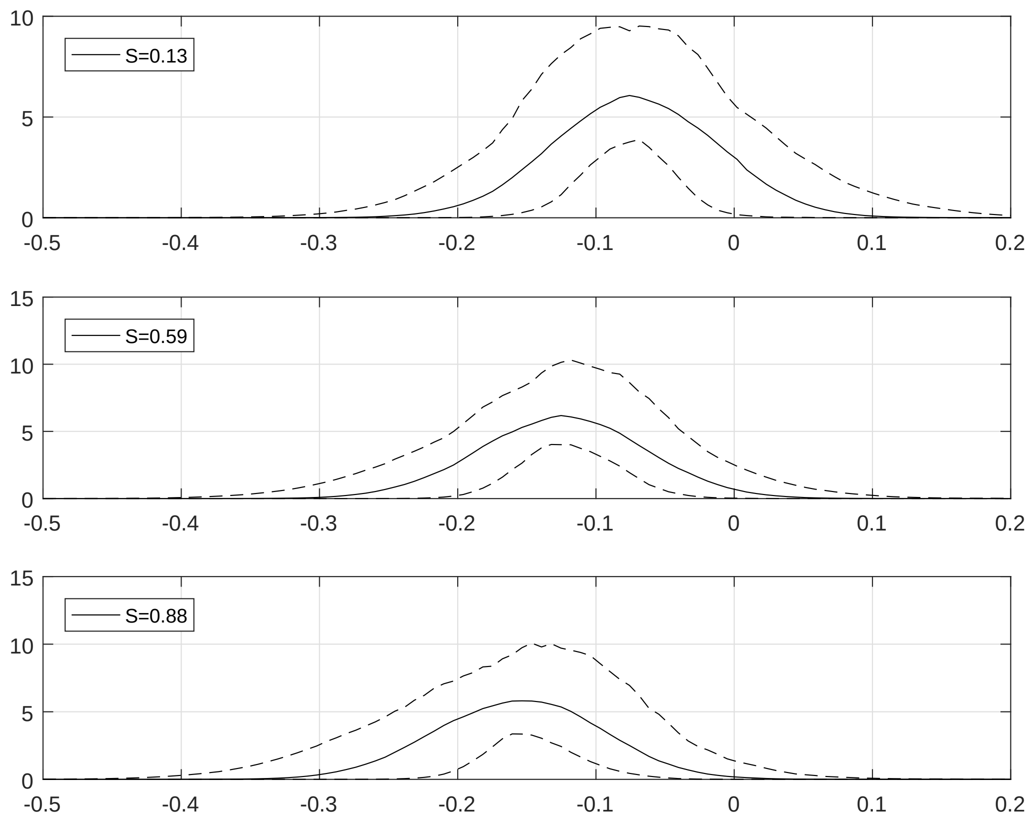

Figure 20: See caption for Fig. 8.

Conditional mean and standard deviation of dollar trade elasticity

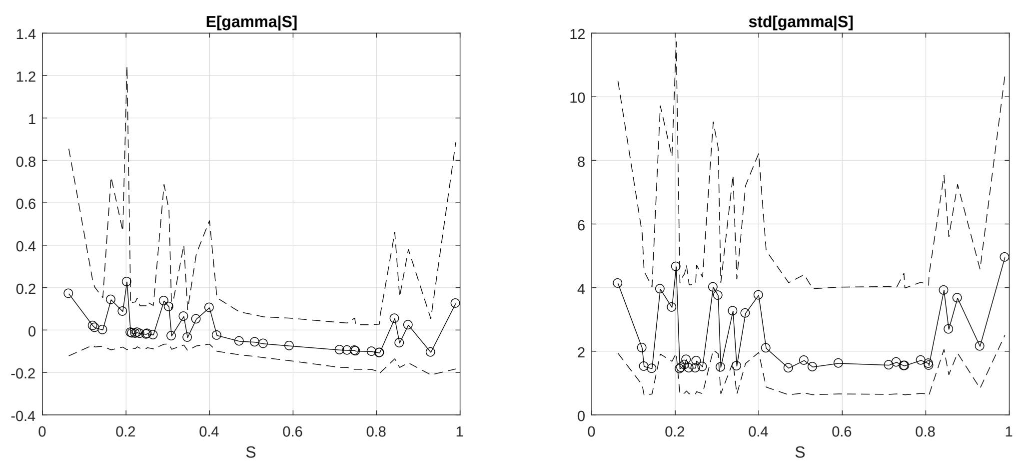

Figure 21: See caption for Fig. 9.

Sample distribution of dollar trade elasticity

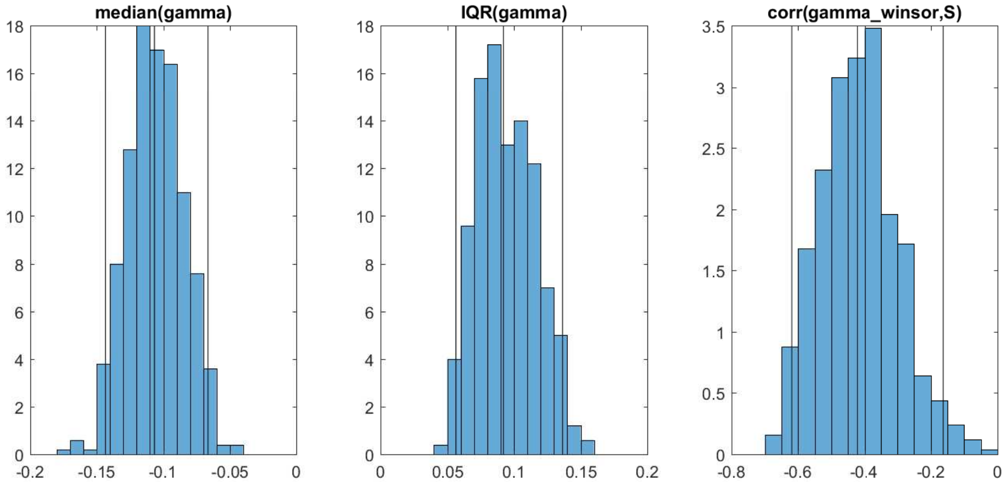

Figure 22: See caption for Fig. 10.

B.3.2 Additional model parameters

For completeness, we now report posterior summaries of the model parameters that are not of primary interest to us.

First we report results for the price pass-through model with $K = 2$. Fig. 23 reports the posterior distribution of the cross-sectionally constant regression coefficients. The results are consistent with the panel regressions in Section 4. In particular, the lagged exchange rate changes are economically insignificant. The posterior for the parameter $\bar{\gamma}$ (the sum of the dollar and bilateral pass-throughs) is concentrated close to 1, indicating near-complete *total* pass-through within a year. Fig. 24 reports the posterior of the mean α and standard deviation τ of the random effects distribution for the dyad-specific effects λ_{ij}, as well as the idiosyncratic standard error σ.

Figs. 25 and 26 provide the same posterior summaries for the trade elasticity model with $K = 4$. Again, these results are consistent with the panel regressions from Section 4.

B.3.3 Robustness to number of mixture components

Here we show that the results in Section 5 are robust to varying the number K of components in the MGLR prior for the cross-sectional distribution of dollar pass-through. Specifically, we here report results for $K = 8$. Figs. 27 and 28 are the $K = 8$ analogues of the price pass-through Figs. 8 and 10 (which had $K = 2$), while Figs. 29 and 30 are the $K = 8$ analogues of the trade elasticity Figs. 20 and 22 (which had $K = 4$). Clearly, the additional mixture components in the $K = 8$ specifications receive very low posterior probability.

Posterior of constant regression coefficients, price pass-through

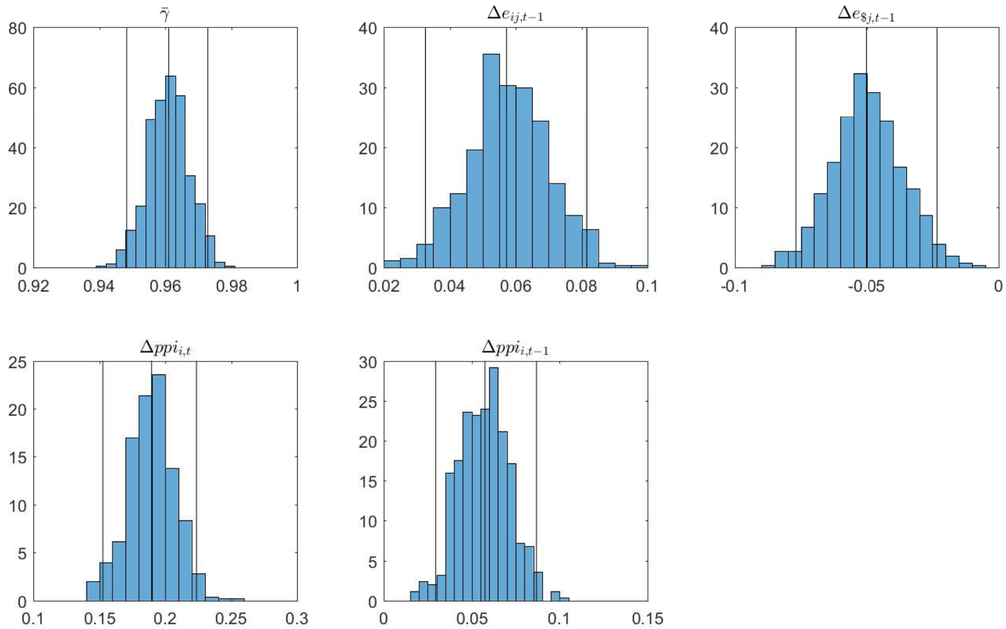

Figure 23: Histogram of posterior draws of elements in θ, the regression coefficients that are assumed constant across dyads. The top left display shows the parameter $\bar{\gamma}$ in Eq. (9). The remaining displays show the coefficients on the indicated exogenous covariates. Vertical lines mark the 2.5, 50, and 97.5 percentiles. For brevity, we do not show the time fixed effects.

Posterior of other parameters, price pass-through

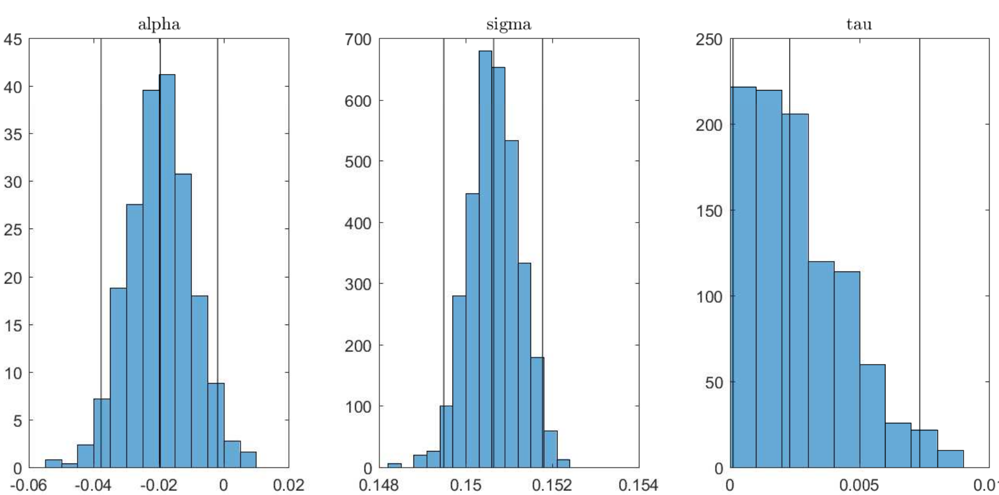

Figure 24: Histogram of posterior draws of α (left), σ (middle), and τ (right). Vertical lines mark the 2.5, 50, and 97.5 percentiles.

Posterior of constant regression coefficients, trade elasticity

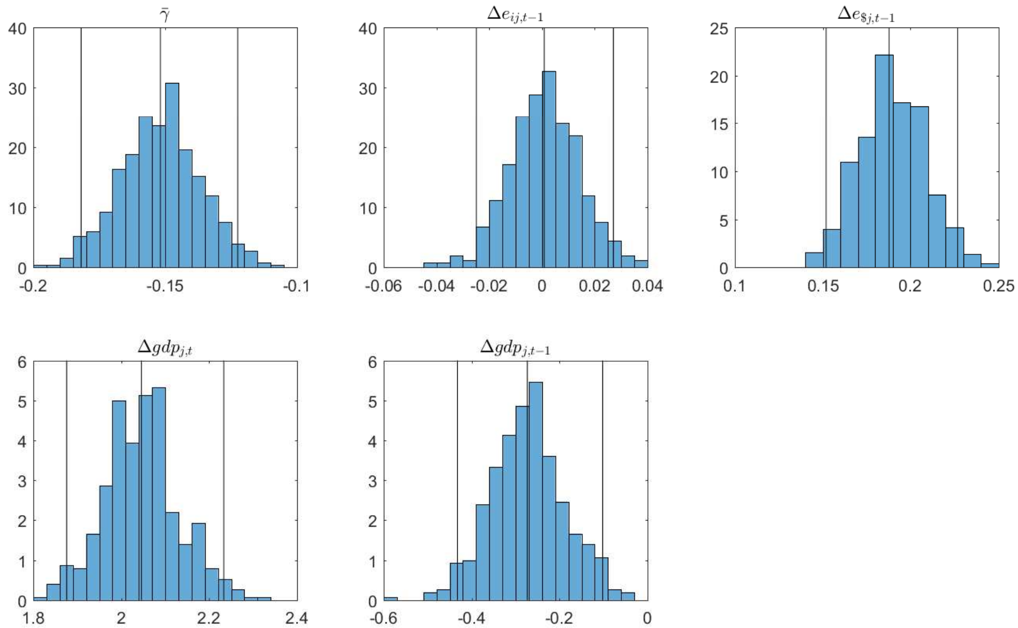

Figure 25: See caption for Fig. 23.

Posterior of other parameters, trade elasticity

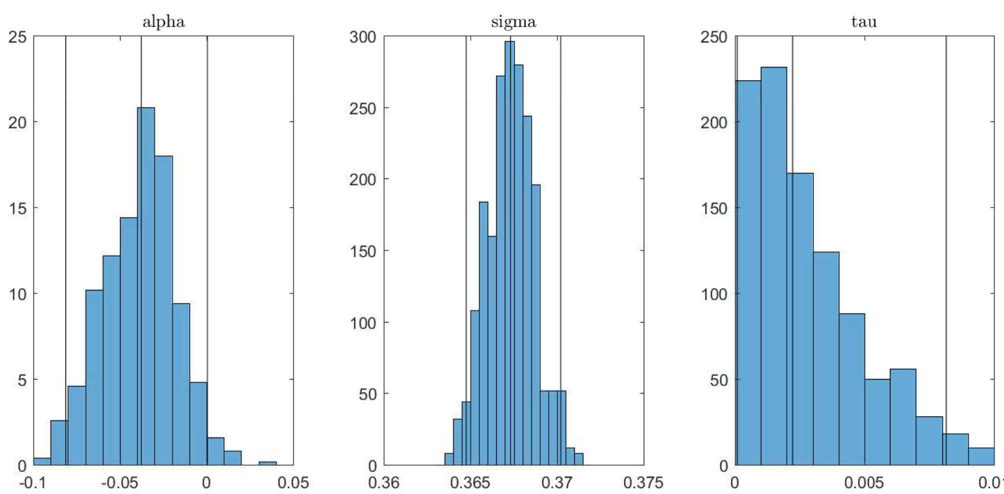

Figure 26: See caption for Fig. 24.

Density of dollar price pass-through given dollar invoicing share, $K = 8$

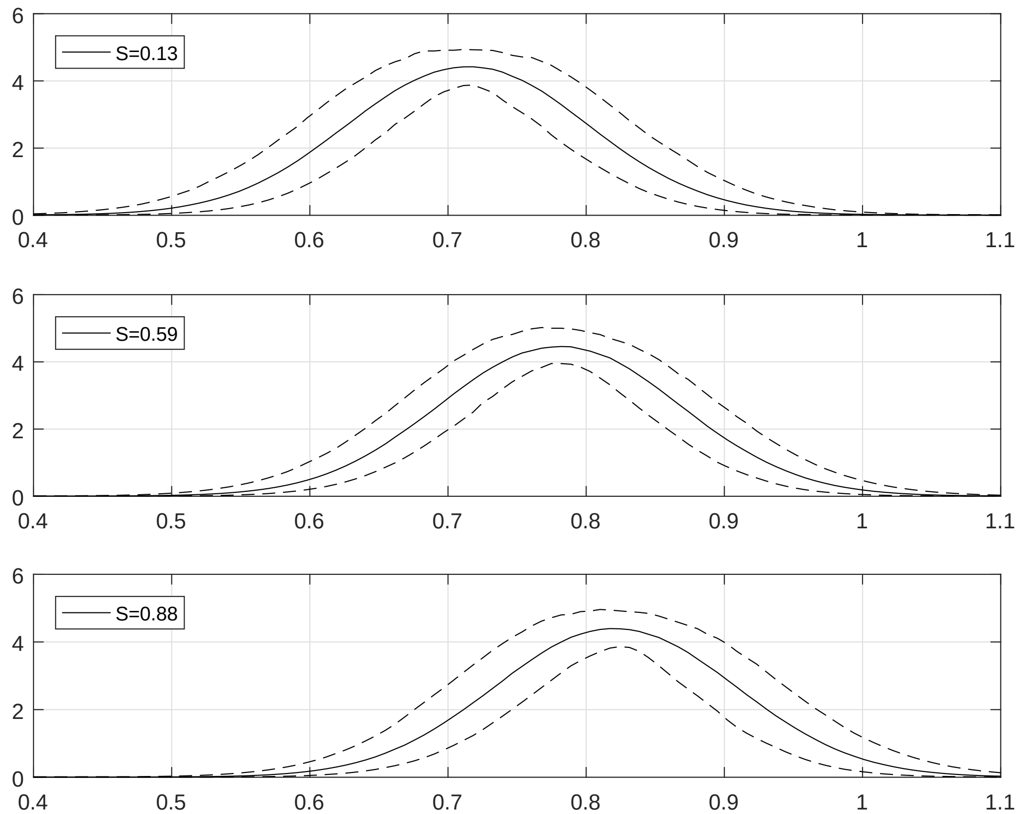

Figure 27: See caption for Fig. 8.

Sample distribution of dollar price pass-through, $K = 8$

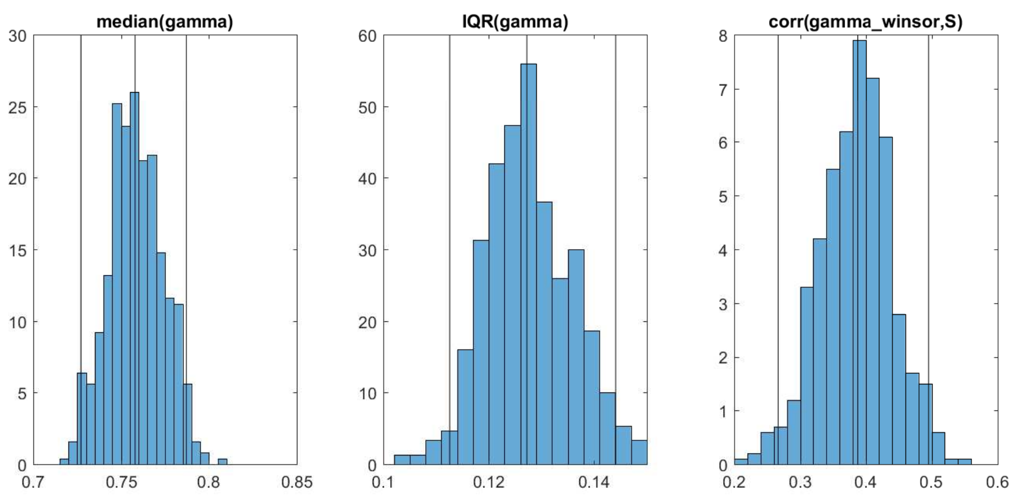

Figure 28: See caption for Fig. 10.

Density of dollar trade elasticity given dollar invoicing share, $K = 8$

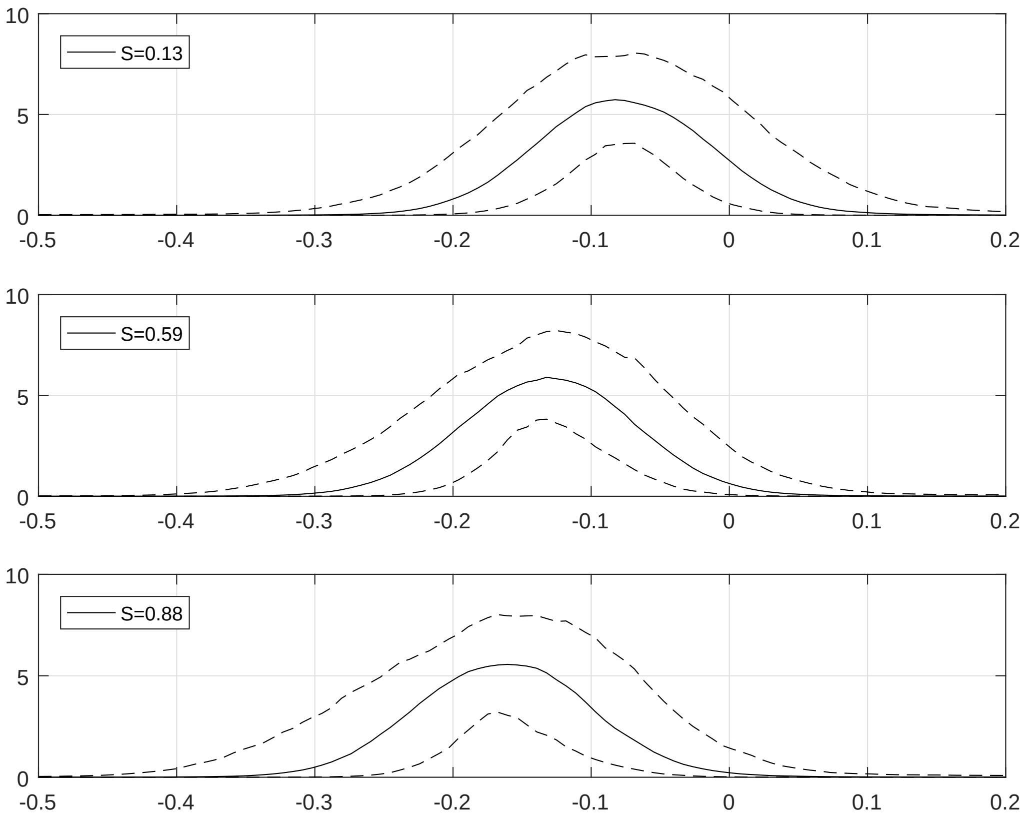

Figure 29: See caption for Fig. 20.

Sample distribution of dollar trade elasticity, $K = 8$

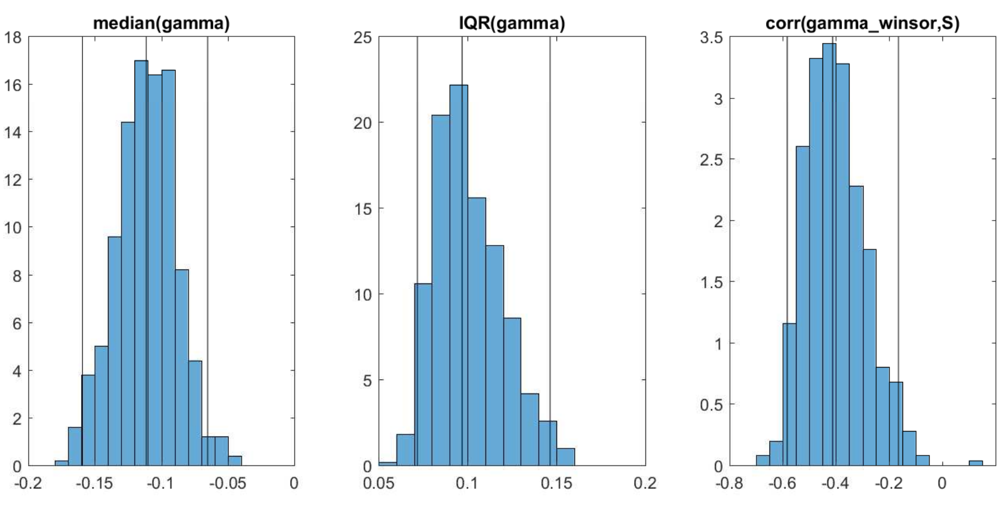

Figure 30: See caption for Fig. 22.

B.3.4 MCMC settings and diagnostics

We execute Stan through Matlab R2016b using MatlabStan 2.7.0.0, which in turn calls CmdStan 2.14.0. For each model specification, we run Stan's No U-Turn Sampler for 2,500 iterations after discarding 1,000 warm-up iterations, storing every 5th draw. The MCMC routine is initialized at parameter values drawn uniformly at random (after the parameters have been transformed to unconstrained support). We use Stan's default settings for adaptively tuning the MCMC routine in the warm-up phase. Our results are completely insensitive to the initialization.

The sampler robustly delivers near-independent draws from the posterior distribution in reasonable time. The stored posterior draws of most model parameters exhibit essentially zero serial correlation after a handful of lags. The only parameters that do not exhibit rapid mixing are those MGLR parameters $\mu_{0,k}, \mu_{1,k}, \omega_k, A_k$ that correspond to mixture components k with low posterior probability $\pi_k(\cdot)$ in model specifications with large K, but these parameters negligibly influence the features of the posterior that we care about. Depending on K and the random initial parameter draw, it takes 2–60 hours to run the MCMC routine for each specification on a personal laptop with a 2.30 GHz processor and 8 GB RAM (no parallel computing is involved). In our experience, it is often sufficient to run the algorithm for 2–4 hours to get a sense of the results.

B.4 Model details

Here we describe in detail the model in Section 6 as well as its calibration.

B.4.1 Households

Each country i is populated with a continuum of symmetric households. In each period household h consumes a non-traded good $C_t(h)$. Each household also sets a wage rate $W_t(h)$ and supplies an individual variety of labor $N_t(h)$ in order to satisfy demand at this wage rate. Households own all domestic firms. The per-period utility function is separable in consumption and labor and given by

$$U(C_{i,t}, N_{i,t}) = \frac{1}{1-\sigma_c} C_{i,t}^{1-\sigma_c} - \frac{\kappa}{1+\varphi} N_{i,t}^{1+\varphi},$$

where $\sigma_c > 0$ is the household's coefficient of relative risk aversion, $\varphi > 0$ is the inverse of the Frisch elasticity of labor supply and κ scales the disutility of labor.

Households solve the following optimization problem,

$$\max_{C_{i,t}, W_{i,t}, B^U_{i,t+1}, B_{i,t+1}(s')} \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t U(C_{i,t}, N_{i,t}),$$

subject to the per-period budget constraint expressed in home currency,

$$P_{i,t}C_{i,t} + \mathcal{E}^U_{i,t}(1+i^U_{i,t})B^U_{i,t} + B_{i,t} = W_{i,t}(h)N_{i,t}(h) + \Pi_{i,t} + \mathcal{E}^U_{i,t}B^U_{i,t+1} + \sum_{s'\in\mathcal{S}} Q_{i,t}(s')B_{i,t+1}(s'),$$

where $P_{i,t}$ is the price index for the final consumption good $C_{i,t}$ that is non-traded. $\Pi_{i,t}$ represents domestic profits that are transfered to households who own the domestic firms. Households also trade a risk-free international bond denominated in dollars that pays a nominal interest rate $i^U_{i,t}$. $B^U_{i,t+1}$ are the dollar holdings of the international bond purchased at time t. $\mathcal{E}^U_{i,t}$ is the dollar exchange rate of country i, that is the price of a dollar in terms of currency i and $\mathcal{E}^U_{U,t} = 1$. Households also have access to a full set of domestic state contingent securities (in their own currency) that are traded domestically and in zero net supply. $Q_{i,t}(s)$ is the period-t price of the security that pays one unit of home currency in period $t+1$ and state $s \in \mathcal{S}$, and $B_{i,t+1}(s)$ are the corresponding holdings. Inter-temporal optimality conditions for U bonds and local bonds are

$$C_{i,t}^{-\sigma_c} = \beta(1+i^U_{i,t})\mathbb{E}_t C_{i,t+1}^{-\sigma_c} \frac{P_{i,t}}{P_{i,t+1}} \frac{\mathcal{E}^U_{i,t+1}}{\mathcal{E}^U_{i,t}},$$

$$C_{i,t}^{-\sigma_c} = \beta(1+i_{i,t})\mathbb{E}_t C_{i,t+1}^{-\sigma_c}\frac{P_{i,t}}{P_{i,t+1}},$$

where $(1+i_{it})$ is the interest rate on i currency bonds.

Households are subject to a Calvo friction when setting wages in i currency: in any given period, they may adjust their wage with probability $1-\delta_w$, and maintain the previous-period nominal wage otherwise. They face a downward sloping demand for the specific variety of labor they supply given by, $N_{i,t}(h) = \left(\frac{W_{i,t}(h)}{W_{i,t}}\right)^{-\vartheta} N_{i,t}$, where $\vartheta > 1$ is the constant elasticity of labor demand and W_t is the aggregate wage rate. The standard optimality condition for wage setting is given by:

$$\mathbb{E}_t \sum_{s=t}^{\infty} \delta_w^{s-t}\Theta_{it,s}N_{is}W_{is}^{\vartheta(1+\varphi)}\left[\frac{\vartheta}{\vartheta-1}\kappa P_{is}C_{is}^{\sigma}N_{is}^{\varphi} - \frac{\bar{W}_{it}(h)^{1+\vartheta\varphi}}{W_{is}^{\vartheta\varphi}}\right] = 0,$$

where $\Theta_{it,s} \equiv \beta^{s-t}\frac{C_{is}^{-\sigma_c}}{C_{it}^{-\sigma_c}}\frac{P_{it}}{P_{is}}$ is the stochastic discount factor between periods t and $s \geq t$ used to discount profits and $\bar{W}_{it}(h)$ is the optimal reset wage in period t.

B.4.2 Non-traded sector

This sector combines all domestic and imported varieties of intermediate goods to produce a good that is consumed as a final good and used as an intermediate input for production. The aggregator $\mathcal{F}$ is implicitly defined by a Kimball (1995) homothetic demand aggregator:

$$\sum_i \frac{1}{|\Omega_j|}\int_{\omega\in\Omega_j}\gamma_j\Upsilon\left(\frac{|\Omega_j|\mathcal{F}_{ji}(\omega)}{\gamma_j\mathcal{F}_i}\right)d\omega = 1. \tag{11}$$

In Eq. (11), $\mathcal{F}_{ji}(\omega)$ represents the demand in country i of variety ω produced by country j, where $j \in \{U, G, R\}$. γ_j is a parameter that is used to captures home bias in j and $|\Omega_j|$ is the measure of varieties consumed in country j. The function Υ is increasing and concave, with $\Upsilon(1) = 1$, $\Upsilon'(\cdot) > 0$ and $\Upsilon''(\cdot) < 0$. This demand structure gives rise to strategic complementarities in price setting resulting in variable mark-ups.

This sector is assumed to be perfectly competitive. Accordingly the break even price index $P_{i,t}$ for the bundle satisfies the expenditure equation

$$P_{i,t}\mathcal{F}_{i,t} = \int_j\int_\Omega P_{ji,t}(\omega)\mathcal{F}_{ji,t}(\omega)d\omega,$$

The demand for any variety given aggregate demand $\mathcal{F}_{i,t}$ is

$$\mathcal{F}_{ji,t}(\omega) = \gamma_j \Psi \left(D_{i,t} \frac{P_{ji,t}(\omega)}{P_{i,t}} \right) \mathcal{F}_{i,t},$$

where $\Psi(\cdot) \equiv \Upsilon'^{-1}(\cdot)$ and $\Psi'(\cdot) < 0$, and $D_{i,t} \equiv \sum_i \int_{\Omega_i} \Upsilon' \left(\frac{|\Omega_i| \mathcal{F}_{ji,t}(\omega)}{\gamma_j \mathcal{F}_{i,t}} \right) \frac{\mathcal{F}_{ji,t}(\omega)}{\mathcal{F}_{i,t}} d\omega$. Define the elasticity of demand $\sigma_{ji,t}(\omega) \equiv -\frac{\partial \log \mathcal{F}_{ji,t}(\omega)}{\partial \log P_{ji,t}(\omega)}$. The log of the mark-up is $\mu_{ji,t}(\omega) \equiv \log \left(\frac{\sigma_{ji,t}}{\sigma_{ji,t}-1} \right)$ and the elasticity of the mark-up is $\Gamma_{ji,t}(\omega) = \frac{\partial \mu_{ji,t}}{\partial \log P_{ji,t}(\omega)}$.

B.4.3 Intermediate goods producers

Each producer manufactures a unique variety ω that is sold both domestically and internationally and purchased by the producers of the non-traded good. The production function uses a combination of labor L_t and intermediate inputs X_t, with a Cobb Douglas production function:

$$Y_{i,t}(\omega) = e^a L_{i,t}^{1-\alpha} X_{i,t}^{\alpha},$$

where α is the constant share of intermediates in production and a is log total factor productivity.

The labor input L_t is a CES aggregator of the individual varieties supplied by each household,

$$L_{i,t} = \left[\int_0^1 L_{i,t}(h)^{(\vartheta-1)/\vartheta} \mathrm{d}h \right]^{\vartheta/(\vartheta-1)}$$

with $\vartheta > 1$.

Markets are assumed to be segmented so firms can set different prices by destination market. Denote $P_{ij,t}^k(\omega)$ the price of a domestic variety ω sold in market j and invoiced in currency k. The per-period profits of the domestic firm producing variety ω are then given by:

$$\Pi_{i,t}(\omega) = \sum_{j,k} \mathcal{E}_{i,t}^k P_{ij,t}^k(\omega) Y_{ij,t}^k(\omega) - \mathcal{MC}_{i,t} Y_{i,t}(\omega),$$

with the convention that $\mathcal{E}_{i,t}^i \equiv 1$, and $Y_{i,t}(\omega) = \sum_{j,k} Y_{ij,t}^k(\omega)$ is the total demand across destination markets and invoicing currencies. Consistent with DCP we assume that for all international sales $k = U$ that is prices are set in dollars. Also all domestic sales are assumed to be priced in the local currency. $\mathcal{MC}_{i,t} = \frac{1}{const.} \cdot \frac{W_{i,t}^{1-\alpha} P_{i,t}^{\alpha}}{e^{a_{i,t}}}$ is the marginal cost of domestic firms in country i, where $const. = (1-\alpha)^{1-\alpha} \alpha^{\alpha}$.

The optimality conditions for hiring labor are given by

$$(1-\alpha)\frac{Y_{i,t}}{L_{i,t}} = \frac{W_{i,t}}{\mathcal{MC}_{i,t}}, \qquad L_t(h) = \left(\frac{W_{i,t}(h)}{W_{i,t}}\right)^{-\vartheta} L_{i,t},$$

with

$$W_{i,t} = \left[\int W_{i,t}(h)^{1-\vartheta} dh\right]^{\frac{1}{1-\vartheta}}.$$

The demand for intermediate inputs is determined by

$$\alpha\frac{Y_{i,t}}{X_{i,t}} = \frac{P_{i,t}}{\mathcal{MC}_{i,t}}.$$

B.4.4 Pricing

Firms choose domestic currency prices at which to sell at home and dollar prices in international markets with prices reset infrequently ala Calvo with probability $1-\delta_p$. The reset price $\bar{P}^k_{ij,t}(\omega)$ of a firm in country i selling in country j and invoicing in currency k satisfies the following optimality condition:

$$\mathbb{E}_t \sum_{s=t}^{\infty} \delta_p^{s-t} \Theta_{t,s} Y^k_{ij,s|t}(\omega)(\sigma^k_{ij,s}(\omega)-1)\left(\mathcal{E}^k_{i,s}\bar{P}^k_{ij,t}(\omega) - \frac{\sigma^k_{ij,s}(\omega)}{\sigma^k_{ij,s}(\omega)-1}\mathcal{MC}_{i,s}\right) = 0,$$

with the convention that $\mathcal{E}^i_{i,t} \equiv 1$. In this expression, $Y^k_{ij,s|t}(\omega)$ is the quantity sold in country j invoiced in currency k at time s by a firm in country i that resets prices at time t and $\sigma^k_{ij,s}(\omega)$ is the elasticity of demand. This expression implies that $\bar{P}^k_{ij,t}(\omega)$ is preset as a markup over expected future marginal costs expressed in currency k, $\mathcal{MC}_{i,s}(\omega)/\mathcal{E}^k_{i,s}$, during the duration of the price. Because of strategic complementarities, the markup over expected future marginal costs is not constant.

B.4.5 Interest Rates

The nominal interest rate in each country is set by its monetary authority that is assumed to follow a Taylor rule with inertia:

$$i_{i,t} - i^* = \rho_m(i_{i,t-1} - i^*) + (1-\rho_m)\left(\phi_M \pi_{i,t} + \phi_Y \tilde{y}_{i,t}\right) + \varepsilon_{i,t}.$$

ϕ_M captures the sensitivity of policy rates to consumer price inflation $\pi_{i,t} = \Delta \ln P_{i,t}$, ϕ_Y measures sensitivity to the output gap $\tilde{y}_{i,t}$ and ρ_m captures the inertia in setting rates. $\varepsilon_{i,t}$ evolves according to an $AR(1)$ process, $\varepsilon_{i,t} = \rho_\varepsilon \varepsilon_{i,t-1} + \varepsilon^m_{i,t}$.

B.4.6 Dollar interest rate

As in Schmitt-Grohé and Uribe (2003), for stationarity purposes in a linearized solution, we assume the dollar interest rate countries $i \neq U$ face is given by

$$i^U_{i,t} = i_{U,t} + \psi(e^{(B^U_{i,t+1}/P_{Ut}) - \bar{B}^U_i} - 1) + \epsilon^U_{i,t},$$

where $\psi > 0$ measures the responsiveness of the dollar rate to the country's net foreign position $B^U_{i,t+1}$. $\bar{B}^U_i$ is the steady state (exogenous) dollar denominated debt (scaled by U price level).[25] The dollar interest rate is an increasing function of the deviation of the aggregate level of debt from the steady state level of debt. Because of the dependence on aggregate debt individual households do not internalize the effect of their borrowing choices on the interest rate.

B.4.7 Market clearing

The goods, labor, and bond markets all clear:

$$\begin{aligned}
Y_{i,t}(\omega) &= \sum_{j \in \{U,G,R\}} \mathcal{F}_{ij}(\omega), \\
N_{i,t} &= L_{i,t}, \\
0 &= \sum_{j \in \{U,G,R\}} B^U_{j,t}, \\
B^k_{i,t} &= 0 \qquad \forall k \notin U.
\end{aligned}$$

B.4.8 Calibration

To simulate the model we use the parameter values listed in Table 14.

[25] This is a standard assumption in the SOE literature to induce stationarity of $B^U_{i,t}$ in a log-linearized environment.

Parameter values for calibrated model

	Parameter	Value
Household Preferences		
Discount factor	β	0.99
Risk aversion	σ_c	2.00
Frisch elasticity of N	φ^{-1}	0.50
Disutility of labor	κ	1.00
Labor demand elasticity	ϑ	4.00
Production		
Intermediate share	α	2/3
(log) Productivity	a	1
Demand		
Elasticity	σ	2.00
Super-elasticity	ϵ	1.00
Home-bias	γ	0.70
Rigidities		
Wage	δ_w	0.85
Price	δ_p	0.75
Monetary Rule		
Inertia	ρ_m	0.50
Inflation sensitivity	ϕ_M	1.5
Output gap sensitivity	ϕ_Y	0.50/4
Shock persistence	ρ_ε	0.50
SS. interest rate	i^*	$(1/\beta) - 1$
Dollar interest rate		
Sensitivity to NFA	ψ	0.001
Steady state NFA	$\bar{B}^U$	0

Table 14: Parameter values for calibrated model in Section 6.