Foreign Official Demand for U.S. Debt and U.S. Yields: Accounting for Common Factors*

Rashad Ahmed† Alessandro Rebucci‡

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Abstract

This paper examines how common factors in foreign yields shape the impact of foreign official purchases of U.S. Treasuries (USTs) on U.S. Treasury yields. We show that the impact of foreign official UST flows on U.S. yields is understated by 30-50% when common factors in non-U.S. sovereign yields are omitted in both reduced-form OLS regressions and VARs of U.S. yields identified through heteroskedasticity. This implies that foreign official UST flows have a substantially larger price impact than suggested by previous estimates. We document that this bias arises because of the pro-cyclical nature of foreign official UST accumulation. Foreign officials buy (sell) USTs against rising (falling) U.S. and foreign yields. Specifically, the common factor in foreign yields co-moves strongly with U.S. yields, with rising foreign yields linked to global economic growth, UST purchases by foreign officials, and UST sales by private investors. This evidence points toward pro-cyclical foreign official flows absorbing counter-cyclical private investor demand for safe assets. Regression estimates of the impact of foreign official flows on U.S. yields increase further after controlling for Federal Reserve large scale asset purchases, suggesting that the effects of foreign official flows are dampened by counter-cyclical Federal Reserve bond purchases similarly to how they are dampened by counter-cyclical private investor flows.

Keywords: Capital flows; Common Factors, Global savings glut; International reserves; Monetary policy; Yield curve
JEL Classifications: E43, E44, F21, F30, G10

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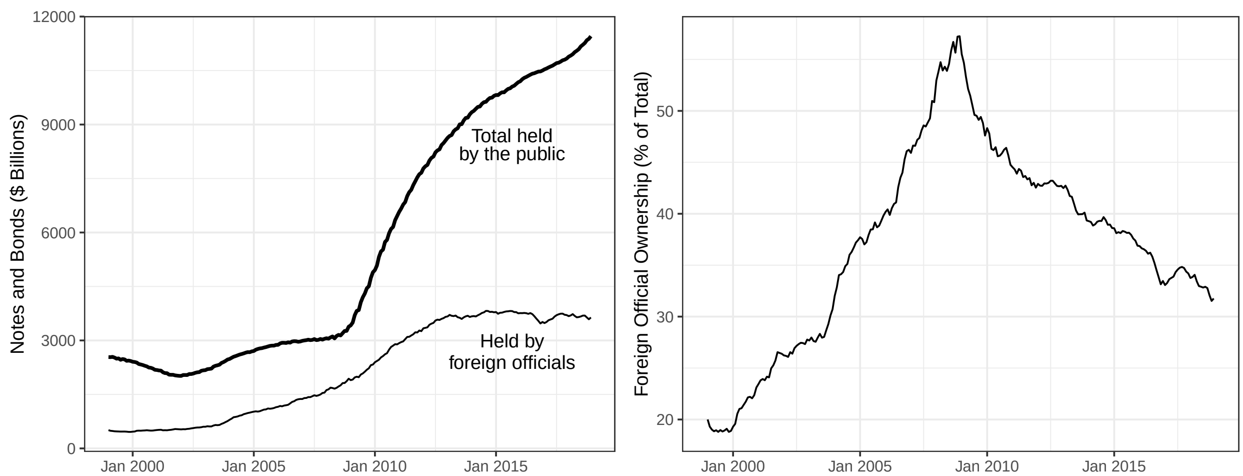
†U.S. Department of the Treasury, Office of the Comptroller of the Currency. E-mail address: rashad.ahmed@occ.treas.gov.

‡Johns Hopkins Carey Business School, CEPR and NBER. E-mail address: arebucci@jhu.edu.

1 Introduction

The 2008 Global Financial Crisis, COVID-19 pandemic, and resulting issuance of central bank swap lines revealed the financial stability risks of concentrated ownership of U.S. Treasuries (USTs) by foreign officials (Figure 1).[1] These episodes not only confirm the inelastic nature of foreign official UST demand [Alfaro et al., 2014; Tabova and Warnock, 2021] but also point toward its distinctly pro-cyclical behavior: accumulation during global expansions and liquidations amid downturns. Unlike foreign officials, private investors exhibit counter-cyclical demand for USTs and global safe assets. As a result, foreign officials buy (sell) USTs when yields are rising (falling) globally. This context motivates a previously overlooked consideration: common cyclical foreign factors jointly shaping and possibly confounding the effects of foreign official UST demand on U.S. Treasury yields.

Figure 1: Foreign Official Ownership of U.S. Treasury Notes and Bonds



Excludes T-bills and includes notes and bonds held by the Federal Reserve (about $2.1 trillion in 2018). Left panel: Total U.S. Treasury notes and bonds (held by the public), along with Treasury notes and bonds held by foreign official institutions. Right panel: Foreign official ownership of notes and bonds as a percentage of total notes and bonds (held by the public). Foreign official holdings based on benchmark-consistent data following Bertaut and Tryon [2007] and Bertaut and Judson [2014].

This paper estimates the impact of foreign official demand for USTs on U.S. interest

[1] Estimates on Treasury security holdings by foreigners from the U.S. Treasury suggest foreign UST sales in the first quarter of 2020 in the range of $300 billion, about half attributed to foreign official institutions. See also Marques et al. [2020], Setser [2020], Aizenman et al. [2021], He et al. [2021], Vissing-Jorgensen [2021].

rates controlling for foreign common yield factors. The extant literature, including the large number of papers listed in Table A.1, omits these factors. Existing estimates vary widely, indicating that a $100 billion foreign official purchase (sale) of USTs can lower (raise) long-term yields by 10 to 90 basis points. Estimates at the upper end of this range would suggest sizable effects, consistent with the 'Global Savings Glut' Hypothesis.[2] But the margin of uncertainty across different studies is wide.

Previous estimates assume that U.S. yields are exclusively determined by domestic factors such as the short rate, inflation, or growth expectations of the United States. Most researchers therefore focus efforts on overcoming the simultaneity problem between foreign UST flows and U.S. yields. The literature typically handles this by separating foreign *official* flows from aggregate foreign flows under the assumption that foreign official demand is inelastic because reserve managers do not maximize risk-adjusted returns. Some studies propose instruments such as FX interventions or trade flows of granular countries to isolate variation in foreign official demand linked to inelastic reserves accumulation. But the exclusion restriction is violated if these instruments are cyclical – and many of them are – a limitation acknowledged by Bernanke et al. [2004].

The presence of global cyclical factors introduces a second source of endogeneity on which the literature has hitherto remained remains silent. U.S. yields respond to domestic factors but are also influenced in a complex way by observed and unobserved global factors, such as current and expected global economic conditions and investor demand for global safe assets. This is evidenced by the remarkably strong co-movement of U.S. yields with yields of other advanced economies [Del Negro et al., 2019]. Moreover, global factors also drive foreign official UST demand because the precautionary and mercantilist motives behind

[2] See Greenspan [2005] on the 'Interest Rate Conundrum', Bernanke [2005], Acharya and Schnabl [2010] and Caballero et al. [2017] on the 'Global Savings Glut'. One Hundred Tenth Congress [2007] and Rogoff [2007] on the 2007 U.S. Congressional hearing on this issue. Caballero and Krishnamurthy [2009] and Krishnamurthy and Vissing-Jorgensen [2012] suggest that aggregate demand for Treasury debt is linked to its special safety and liquidity properties. Bernanke et al. [2011], Du et al. [2018], Krishnamurthy and Lustig [2019] and Jiang et al. [2021] argue that these properties are particularly important attractors of foreign demand for USTs which can drive U.S. interest rates away from fundamentally-justified levels.

foreign official UST accumulation tend to be cyclical. For instance, global economic booms are accompanied by rising safe asset yields, growing export demand, and capital inflows. Foreign central banks respond by accumulating international reserves to stem appreciating exchange rates or to build precautionary buffers. Conversely, economic downturns reverse these dynamics and induce foreign officials to sell reserves for liquidity purposes or in order to stabilize the currency. So while foreign official demand for USTs is inelastic to an extent, it also depend on the state of the global economy. As a result, estimates that account for simultaneity may remain biased because the endogeneity arising from the presence of (omitted) global factors is left unaddressed.

This paper extends this literature by being the first to consider how global factors shape the interaction between foreign official demand for U.S. debt and U.S. interest rates. Approximating global factors using average foreign long-term and short-term international bond yields as a benchmark, the central result of this study suggests that failing to condition on common factors in foreign yields leads to a significant understatement – often by 50% or more – of the impact of foreign official demand for USTs on U.S. long-term yields. This bias arises out of foreign officials liquidating USTs amid cyclical downturns, precisely when weaker global growth and investor flight to safety exert downward pressure on U.S. and foreign yields (when demand for global safe assets is highest [Lustig et al., 2014]). Similarly, foreign officials tend to accumulate USTs during cyclical expansions amid higher growth, risk appetite and inflation, when investor demand for global safe assets weakens.

Adding further to this view, we show that failing to control for Federal Reserve large scale asset purchases (LSAP) also leads to an understatement of the impact of foreign official UST flows on U.S. yields. Fed LSAP programs are counter-cyclical, typically executed amid global crises. Therefore, the Federal Reserve, like U.S. households and mutual funds, buy U.S. Treasuries as foreign officials are selling them.

For policymakers and researchers concerned with financial stability, a practical implication becomes apparent. *Pro-cyclical* foreign official demand can be an important source of

Treasury market liquidity because it absorbs *counter-cyclical* private investor demand.[3] At the same time, sudden UST liquidations or purchases of the black-swan type may have a substantially larger impact on U.S. financial conditions than implied by the existing literature. Finally, these results suggest that the effectiveness of government asset purchases also depend on the market conditions in which they occur.

The sample under consideration is the period January 1999 to December 2018 or February 2021 depending on the flows data applied. The period intentionally spans two decades characterized by rapid global economic and financial integration, but longer sample periods are also examined. Foreign factors associated with U.S. Treasury yields are recovered from a panel of short-term and long-term government bond yields of 19 non-U.S. advanced economies. Not only are foreign yield factors significantly associated with U.S. yields, they are also positively associated with world economic activity, positively associated with accumulation of USTs by foreign officials, and negatively associated with UST demand from U.S. households, hedge funds, and mutual fund investors. Taken together, these patterns suggest that foreign yield factors indeed reflect global cyclical forces, and that foreign officials buy and sell USTs 'going against the heard', as global yields rise and fall, respectively.

We report both OLS and identified estimates of the impact of foreign official UST demand on U.S. yields. We start by augmenting the regression framework of Warnock and Warnock [2009] with foreign common yield factors. While this specification does not control for simultaneity, it provides a transparent benchmark. A \$100 billion foreign official sale of USTs is associated with a rise in U.S. 10-year yields and term premia of about 60.5 and 42 basis points, respectively (flow effect), which is substantially larger than estimates that omit global controls (19 and 13.6 basis points, respectively).[4] Similarly, a \$100 billion change in the *stock* of foreign official UST holdings is associated with an impact of 5.8 basis points

[3] More generally, this view is consistent with Jiang et al. [2011] who document the importance of two-sided order flow in explaining price movements in Treasury markets.

[4] For ease of interpretation, we refer to the effects of foreign official UST flows on yields in terms of either purchases or sales throughout this paper but these effects are symmetric. In fact, we do not find evidence supporting asymmetric effects or foreign official UST purchases versus sales.

on U.S. 10-year yields compared to 3 basis points in the absence of global factors (stock effect). Conditioning on Fed LSAP shocks of Swanson [2021] further increases the impact of foreign official UST flows on 10-year U.S. yields. This result is again consistent with the confounding role of cyclical factors, since the Federal Reserve tends to announce bond-buying programs counter-cyclically, during economic downturns when foreign officials tend to sell USTs. These results are robust to a number of checks: addressing two-way feedback between U.S. and foreign yields, alternative econometric specifications, using raw Treasury International Capital (TIC) flows data rather than the benchmark-consistent foreign official flows of Bertaut and Tryon [2007] and Bertaut and Judson [2014], different sample periods, and incorporating additional global factors.

We then extend the baseline regression to a structural vector autoregression (VAR) for short, medium and long-term U.S. yields. A UST purchase shock is identified through heteroskedasticity [Rigobon, 2003; Brunnermeier et al., 2021] exploiting a well documented regime change in the pattern of global capital flows. Specifically, we exploit the change in the volatility of capital flows and the dollar factor in exchange rates that occurred around the time of the 2008 global financial crisis. The identified estimated impulse responses show that a $100 billion foreign official UST sale or purchase is causes a change in 5-year yields of 27.5 basis points and 10-year yields of about 35.75 basis points, compared to effects roughly half this size when omitting global factors. The impact increases with maturity, consistent with foreign official flows affecting the U.S. term premium.

For robustness, we also consider other identification strategies to validate our findings. First, we consider a Cholesky decomposition where foreign official flows are ordered first in the VAR, consistent with the OLS regression specification, assuming that foreign official demand for USTs is inelastic and hence exogenous after accounting for global factors. We also try ordering foreign official flows last to more explicitly guard against simultaneity. IRFs under both approaches are qualitatively consistent with the benchmark identified through heteroskedasticity, but less clearly identified. We also show that the effects of foreign private flows are much weaker than those of foreign official flows. Consistent with the relative inelasticity of foreign official demand, this suggests that confounding global factors may be a more important source of endogeneity than simultaneity. Finally, in the spirit of instrumental variables approaches [Beltran et al., 2013], we exploit China-specific demand in a two-pronged identification strategy that tries to deal with both global factors and simultaneity jointly. Specifically, *lagged* variation in Chinese Renminbi volatility is used to capture official UST demand linked to China's exchange rate policies. These extensions confirm the baseline results and suggest that U.S. yields are sensitive to official demand specifically linked to China.

The rest of the paper is organized as follows. Section 2 constructs foreign common yield factors and presents stylized facts on the cyclical nature of foreign and domestic UST demand and U.S. yields. Section 3 presents the baseline OLS regression and the attendant results. Section 4 sets up our structural VAR identified through heteroskedasticity and reports the main result of the paper. Section 5 explores the potentially confounding role of Federal Reserve bond purchases, and compares estimates of foreign official demand on U.S. yields with estimates from the Quantitative Easing literature. Section 6 reports the main robustness checks on the baseline regression and alternative identification schemes for the VAR. Section 7 concludes. The Appendix provides additional detail on the data, including sources and construction, evidence on the role of global factors in the joint determination of U.S. yields and foreign official UST flows, robustness of global yield factor construction to alternative weighting schemes, alternative model specifications, and several additional results.

2 Common Factors in Foreign Bond Yields, Official Purchases of USTs, and U.S. Yields

Evidence of common factors driving co-movements in government bond yields and neutral rates continues to grow.[5] In the context of foreign demand for USTs, Rachel and Smith [2015], Del Negro et al. [2019] and Ferreira and Shousha [2020] find that rising demand for safe assets is an important factor driving global interest rates, as are global demographic transitions and productivity trends. This has an important implication: movements and trends in U.S. interest rates are part of a global phenomenon.

The presence of global economic forces, though difficult to measure, implies the joint determination of foreign UST demand and U.S. bond yields. Foreign official demand for USTs rises amid global expansions as official institutions accumulate international reserves under mercantilist, precautionary, or exchange rate smoothing motives. At the same time, U.S. yields rise amid global expansions since the U.S. and global economy are interconnected. Therefore, to consistently estimate the impact of foreign demand for USTs on U.S. interest rates and limit omitted variable bias, these foreign factors must be controlled for. As a simple illustration, suppose long-term U.S. yields depend linearly on both domestic and foreign factors along with foreign official UST flows:

$$y^{10Y}_{us,t} = \mathbf{a_1'} F_{us,t} + a_2 \Delta FO_t + \mathbf{a_3'} F^G_t + e_{us,t}, \tag{1}$$

where $y^{10Y}_{us,t}$ is the U.S. 10-year Treasury yield, $F_{us,t}$ represents a set of domestic covariates

[5] See for example, Ilmanen [1995], Sutton [2000], Diebold et al. [2008], Obstfeld [2020], Kim and Ochoa [2021]). Strong co-movement is especially noticeable in long-term bonds across advanced economies since they are not only driven by domestic monetary policy but often contain a sizable term premium [Hellerstein, 2011; Dahlquist and Hasseltoft, 2013]. The presence of common global forces implies that U.S. yields are not merely determined by domestic factors but that interest rates are determined in a global equilibrium [Clarida, 2019]. International factor structure can manifest under globally integrated financial markets [Miranda-Agrippino and Rey, 2020] which lend to no arbitrage in the long-run. Alternatively, global factor structure can rise out of correlated domestic fundamentals across countries, perhaps through deeper trade integration leading to co-movement in inflation [Ciccarelli and Mojon, 2010; Jotikasthira et al., 2015; Byrne et al., 2019], productivity and real activity [Rachel and Smith, 2015; Bekaert and Ermolov, 2021], or common monetary regimes [Borio et al., 2019].

such as the short-term rate, economic fundamentals, etc., ΔFO_t are net foreign official purchases of USTs which can be interpreted as an observable foreign factor, F_t^G is composed of other, possibly *unobserved* foreign factors, and $e_{us,t}$ is an i.i.d. error term. These foreign unobserved factors capture a host of interconnected forces affecting U.S. yields which are likely correlated with ΔFO_t, such as current and expected global macro-financial conditions and investor demand for safe assets. The goal is to consistently estimate a_2, the effect of foreign official net purchases of USTs on 10-year U.S. yields. Previous studies estimate a_2 in (1) under the assumption that foreign official demand is inelastic, but they exclude other foreign factors, F_t^G. So long as F_t^G is a determinant of $y_{us,t}^{10Y}$ and correlated with ΔFO_t, excluding F_t^G biases the effect of foreign official flows on U.S. yields, a_2, even in the absence of simultaneity. More specifically, omitting the role of foreign factors which positively covary with U.S. yields would bias the estimated effect of foreign official UST demand on U.S. yields upward. If the true impact of greater foreign UST demand on Treasury yields is negative, ignoring other foreign factors would bias the estimated impact from negative toward zero or positive since 1) the relationship between U.S. yields and foreign official UST demand is negative, 2) U.S. yields are positively associated with foreign economic factors, and 3) the association between foreign economic factors and foreign official UST demand is positive. Intuitively, this arises from the fact that foreign officials sell USTs during periods when U.S. and foreign yields are falling (i.e. deteriorating economic conditions; private investor flight-to-safety) and *vice versa*.

Empirically, the first challenge is to reasonably approximate the foreign component of U.S. yield variation, F_t^G. As mentioned, there could be multiple global drivers, observed and unobserved, coincident and forward-looking. Motivated by the literature on global bond yield factors and existing work on cross-sectional dependence [Pesaran, 2006; Cesa-Bianchi et al., 2020], we construct foreign factors (F_t^G) using weighted averages of the cross-section of international long and short-term government bond yields. The key idea is that bond yield co-movement captures the foreign component of U.S. yield variation. We consider advanced economy government bonds as they tend to exhibit low credit risk and the strongest co-movement with U.S. Treasury yields. Several advanced economy bonds are also perceived as safe assets and considered reasonable U.S. Treasury substitutes (e.g. German, Japanese, U.K., Swiss bonds).[6] Unlike emerging market yields, advanced economy yields are more likely to reflect counter-cyclical global investor demand for safe assets, a critical source of endogeneity we wish to control for.[7] Moreover, the common component of advanced economy bond yields has been shown to reflect current and expected world economic conditions, which not only shape foreign official UST demand, but are also particularly synchronized with U.S. economic conditions (see Figure A.1).

In constructing foreign factors from yields, it is important to note the key assumption that no underlying country bears a disproportionally large weight. Realistically, this unlikely the case for our main country of interest, the United States. Rather, it is likely that causality runs both ways among advanced economies interest rates: advanced economy yields influence U.S. yields, and *vice versa* [Kim and Ochoa, 2021].[8] Alternatively, one could perhaps construct additional foreign economic factors from measures of real economic activity and inflation across countries, but economic aggregates tend to update with a lag and do not incorporate forward-looking information reflected in asset prices like U.S. Treasuries. There could also be additional foreign factors aside from the bond yield factors that are correlated with both U.S. yields and foreign official UST demand. For robustness, we later explore the relevance of several additional foreign and global factors.

[6]Despite substitutability, foreign demand for USTs should disproportionally affect U.S. yields compared to similar non-U.S. yields, motivating the examination of U.S. yields relative to global advanced economy yields.

[7]Save Japan and Switzerland, most large official UST holders are emerging markets.

Using emerging market yields to construct the global factor component of U.S. Treasury yields would inappropriate for several reasons. In fact, it is in part the foreign official demand associated with idiosyncratic emerging market shocks *uncorrelated* with global investor demand and economic conditions we wish to identify rather than partial out. There are also model specification problems. Emerging market yields have sizable credit risk components [Du and Schreger, 2016], and the model assumption implicit in (1) that F_t^G determines U.S. yields would very likely be violated: U.S. and broad advanced economy conditions affect emerging markets but the reverse is less likely. That said, including emerging market yields are considered in a robustness check.

[8]Although disentangling these two-way feedback effects from each other is difficult in practice, we later explore various robustness checks that try to deal with this problem.

2.1 Estimating the Common Factors in Foreign Yields

We consider monthly 3-month and 10-year government bond yields from the U.S. and 19 advanced economies: Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, The United Kingdom, Ireland, Italy, Japan, Luxembourg, Netherlands, Norway, New Zealand, and Sweden from January 1999 to February 2021. Section A2 in the Appendix presents evidence of a common factor governing these advanced economy bond yields and also U.S. bond yields.

Several approaches can be taken to construct the common long-term and short-term bond yield factors. The estimated first principal component (PC) recovered from the panel of long-term and short-term yields is one approach. However, PC analysis relies on information from the full sample such that the estimated factor value for month t depends on future data sampled at time $t + h$. This might not be desirable especially in a forecasting settings. Alternatively, an approach which doesn't suffer from this limitation is taking the cross-section average (CSA) of yields in each period, as we do in Section A2 for testing factor strength (see also for example, Cesa-Bianchi et al. [2020]). A second advantage of CSA over Principal Components Analysis (PCA) is that for consistent estimation of the factor, CSA only requires large N asymptotics while PCA requires both large N and large T. Practically speaking, the difference between factor estimates using the PCA approach and the CSA approach boils down to the weights assigned to each component. The PCA approach estimates weights for each component, while cross-section averaging assigns equal weights to each individual country yield. In both cases, weights do not vary over time. However, beside statistical approaches, weights can be economically motivated depending on the context, e.g. based on GDP shares or relative supplies of aggregate public debt. Weights based on economic rationale have two favorable features. First, because the weights are not estimated there is less estimation uncertainty than the PC approach for recovering unobserved factors[9], and

[9]This is particularly important when the estimated factor is used in a later stage as an input in regression analysis as we do in this study (known to introduce the generated regressor problem).

second, weights can vary over time – the latter point aligning more closely with reality. Like Hellerstein [2011], we therefore estimate long and short-term bond yield factors as the GDP-weighted averages of long-term and short-term bond yields (respectively) across countries. Specifically, we use lagged GDP denominated in U.S. Dollars, and consider the 19 non-U.S. advanced economy government bond yields to construct the global bond yield factors:

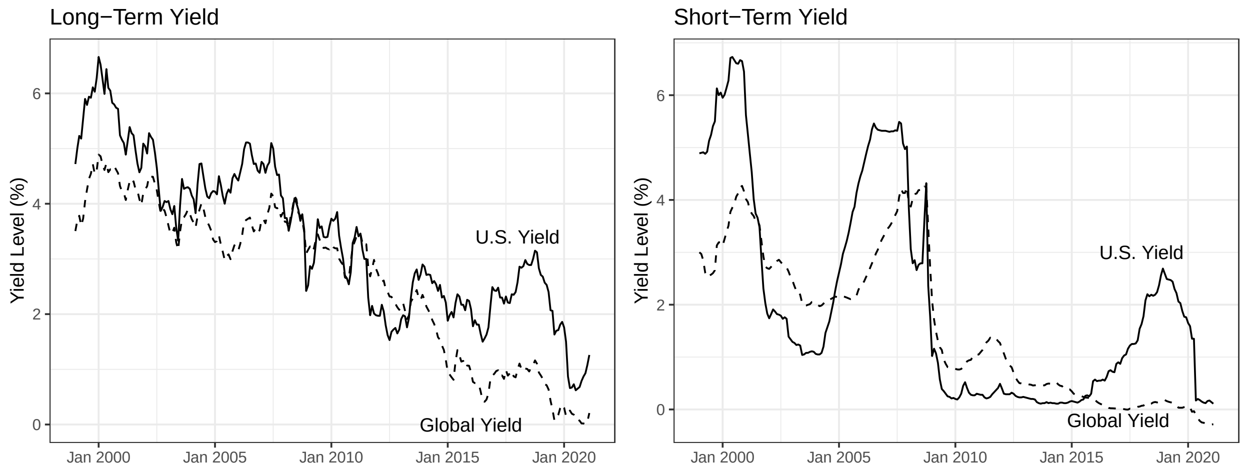
$$\mathcal{Y}^{10Y}_{g,t} = \sum_{i \notin us} w_{i,t-1} y^{10Y}_{i,t}, \quad \mathcal{Y}^{3M}_{g,t} = \sum_{i \notin us} w_{i,t-1} y^{3M}_{i,t}. \tag{2}$$

U.S. yields are omitted here because they are the dependent variable of interest, and the common factors will be used as explanatory variables later in the analysis. The GDP weight for country i in a given month t is computed as the GDP share of that country divided by the total month t GDP of the 19 non-U.S. countries in that period: $w_{i,t} = GDP_{i,t} / \sum_{i \notin us} GDP_{i,t}$.[10] GDP weights are intuitive in that larger countries, based on economic activity, will be more influential in determining variation in the foreign yield factor, which aims to proxy for current and forward-looking economic forces. Alternatively, weights reflecting differences in the supply of national public debt stock across countries could allow for a more market-based approach. For instance, Japan's U.S. Dollar value of public debt is high even relative to GDP, and a debt-based weighting scheme would account for Japan's large market size. Section A3 of the Appendix shows that estimates of the global yield factors are not sensitive to alternative weight schemes by comparing factors estimated under PCA, CSA, GDP, and debt based weights.

Figure 2 traces 10-year and 3-month common foreign yield factors (GDP-weighted) along with U.S. interest rates. Changes in global yields are highly correlated with changes in U.S. yields (0.88 for 10Y yields, 0.77 for 3M yields), but we also see some prolonged periods where U.S. interest rates deviate from the non-U.S. weighted average of foreign yields. Most

[10] Data on quarterly nominal GDP (expenditure approach) in U.S. Dollars is taken from the OECD. To convert the quarterly series to monthly, monthly observations between the quarterly observations (January, April, July, October) are imputed with the most recent previously recorded GDP value. Further detail can be found in Section A1 of the Appendix.

Figure 2: U.S. and Foreign (non-U.S.) Interest Rates



Solid lines are the U.S. 10-year (left panel) and 3-month (right panel) yields and dashed lines are the global 10-year (left panel) and 3-month (right panel) yield factors. Global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2.

recently when the U.S. embarked on a monetary tightening cycle in 2015, both long and short-term U.S. rates rose, substantially diverging from global rates.

Since the start of the sample period in 1999, both short-term and long-term U.S. yields seem to systematically diverge from global yields on three occasions preceding U.S. recessions: in the early 2000s amid the 'dot-com' boom, prior to the 2008 Global Financial crisis, and again prior to the 2020 pandemic recession. All three episodes coincided with Fed monetary tightening cycles. Each period was followed by a U.S. or global recession with U.S. yields appearing to revert down to global yields, but also some degree of global yields reverting to U.S. yields.

In Section A2.1 of the Appendix, we show that these common factors can explain variation in U.S. 10-year and 3-month yields beyond what is explained by traditional domestic term structure factors (level, slope, curvature), suggesting an important role for foreign common in the pricing of U.S. interest rates. The next section focuses on the role of these foreign common yield factors in reflecting broad economic conditions and shaping demand for USTs.

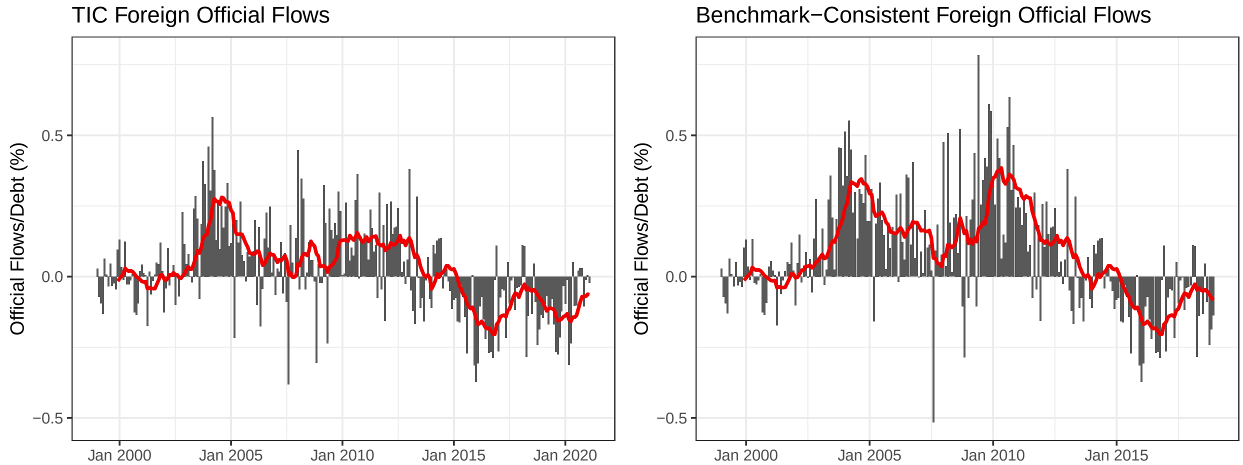
2.2 Comovement among Demand for Treasuries, U.S. Interest Rates, and Common Foreign Factors

The foreign common yield factor estimated in the previous section is strongly associated with U.S. yields, but does it reflect global cyclical conditions and if so, does it shape foreign official demand for USTs? This potential source of endogeneity, that has received little attention in the literature and is typically omitted from the analysis, can jointly influence U.S. yields and foreign (and domestic) demand for USTs. For instance, optimistic global growth expectations can result in broadly higher interest rates globally and in the U.S., and at the same time increased official demand for USTs because greater export demand and/or capital inflows prompt foreign officials to accumulate international reserves, especially in countries where monetary policy limits exchange rate flexibility. At the same time, private investors are likely to rotate out of USTs as risk appetite increases. As a result, episodes of global financial distress and economic downturns, interest rates fall in other advanced economies and in the U.S., as private investors seek safety even while foreign officials, mostly emerging markets, tend to liquidate international reserves.[11]

In Section A4 of the Appendix, we present explicit evidence consistent with this view. We show that advanced economy yields are indeed reflective of global economic conditions. The foreign yield factor is significantly and positively associated with current and future industrial production growth rates in the U.S., advanced economies (ex. U.S.), and emerging market economies (Figure A.1). Figure A.2 shows that foreign yields are significantly and positively associated with current and future demand for USTs by foreign officials (pro-cyclical UST demand) but not with foreign private UST demand. By contrast, movements in foreign yields are significantly and negatively associated with current and future demand for USTs by domestic U.S. households (includes hedge funds) and U.S. mutual funds, reflecting a

[11] Pihlman and van der Hoorn [2010], Dominguez et al. [2012], Bussière et al. [2015], Avdjiev et al. [2017], Jones [2018], Schanz [2019], Ahmed [2020] all document evidence of cyclical reserves management driven by global shocks, while Marques et al. [2020], Setser [2020], He et al. [2021], Vissing-Jorgensen [2021] specifically report large sales of USTs during crises.

Figure 3: Monthly Foreign Official Net UST Purchases of U.S. Treasury Notes and Bonds as a Percentage of U.S. Marketable Debt Outstanding



Bars are monthly purchases/sales. Line is the 12-month rolling average. U.S. marketable debt outstanding is lagged 12 months. Left panel is the TIC-reported flows data and right panel are adjusted benchmark-consistent flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014].

source of counter-cyclical UST demand. In addition, the foreign term spread, measured as the difference between the 10-year and 3-month foreign yield factors, is significantly positively correlated with both net foreign official purchases and with the U.S. term spread (Figure A.3). Unlike foreign officials who accumulate USTs pro-cyclically, investor risk appetite appears to be an important driver of counter-cyclical UST demand from private investors, who buy USTs when foreign officials sell them.

Despite the positive correlation between the common factor in foreign yields and foreign official UST flows, foreign official flows are *negatively* correlated with the U.S.-global 10-year interest rate differential (Figure A.4). While crude, netting out the foreign yield factor may partial out the foreign component of U.S. yields, thereby presenting cursory evidence suggesting that greater foreign official UST demand pushes down U.S. yields relative to global yields. On the other hand, this negative relationship may also point toward causality in the other direction, such that U.S. monetary policy divergence from the rest of the world inducing foreign currency depreciation, prompting reserve managers to intervene and sell their UST holdings.

3 Controlling for Common Foreign Yield Factors

To provide estimates of the impact of foreign official demand for USTs, we start by building a baseline model augmenting Warnock and Warnock [2009] with our foreign common yield factors estimated with OLS. The level of the 10Y U.S. interest rate is regressed on domestic covariates along with foreign official UST flows and common long-term and short-term yield factors. While this specification does not deal with endogeneity, simply assuming that official foreign purchases of USTs is exogenous, it provides a benchmark to compare estimates against those from previous studies. We then address identification in a VAR setting by exploiting regime shifts in the variance of the time series of interest in Section 4.

3.1 Model Specification

The sample period considered is January 1999 to February 2021 using monthly raw TIC data, and January 1999 to December 2018 using the benchmark-consistent flow data. Focusing the sample starting in the early 2000s aligns well with economic era of rapid globalization and international financial development giving rise to the 'Global Savings Glut', though we also examine longer sample periods in a series of robustness checks. Following the previous literature, we use foreign official flows instead of aggregate or foreign private flows is meant to address the simultaneity between U.S. yields and capital flows because foreign official demand tends to be inelastic.[12] The baseline regression model is specified as follows:

$$y^{10Y}_{us,t} = \phi_0 y^{10Y}_{us,t-1} + \phi_1 y^{3M}_{us,t} + \phi_2 \mathcal{Y}^{3M}_{g,t} + \phi_3 \mathcal{Y}^{10Y}_{g,t} + \sum_{l=0}^{L} \theta_l \Delta FO_{t-l} + \boldsymbol{\beta}'\mathbf{X}_t + \epsilon_{us,t}, \tag{3}$$

$$\mathbf{X}_t = [\mathbf{1}, t, \Delta GDP^{E[t+1]}_t, \pi^{E[t+1]}_t, \pi^{E[t+10]}_t, VIX_t, surplus_t],$$

[12] Foreign official UST demand is thought to be driven by the same precautionary [Aizenman and Lee, 2007; Kilian et al., 2009; Jeanne and Ranciere, 2011], mercantalist [Korinek and Serven, 2016], and/or exchange rate smoothing [Obstfeld et al., 2010; Benigno and Fornaro, 2012; Levy-Yeyati et al., 2013; Ito and McCauley, 2020] motives.

where $y_{us,t}^{10Y}$ and $y_{us,t}^{3M}$ are U.S. long and short-term Treasury yields corresponding to the 10-year and 3-month maturity, respectively. Short-term and long-term foreign yield factors are denoted $\mathcal{Y}_{g,t}^{3M}$ and $\mathcal{Y}_{g,t}^{10Y}$, respectively. Foreign official purchases and sales of U.S. Treasury bonds and notes are given by ΔFO_t and are scaled by 12-month lagged U.S. marketable debt outstanding. U.S. marketable debt outstanding is calculated as total U.S. public debt less public debt held by the Federal Reserve banks.[13] Note that for interpretation purposes, the literature typically uses regression estimates to back out the effect of a \$100 billion purchase or sale (rather than a percentage point flow or sale) because the scale of the denominator affects the regression coefficient size. As a robustness check, flows are also scaled by GDP as done in Warnock and Warnock [2009].

The specification includes several domestic controls: Expected 1-year real GDP growth ($\Delta GDP_t^{E[t+1]}$), expected 1-year and 10-year inflation ($\pi_t^{E[t+1]}$ and $\pi_t^{E[t+10]}$), the CBOE VIX index which captures risk appetite (VIX_t)[14], the structural budget surplus or deficit ($surplus_t$) as a percent of GDP along with an intercept term and linear time trend ($\mathbf{1}$ and t). GDP growth, inflation expectations and budget surplus data are originally quarterly frequency data, which we interpolate linearly to monthly frequency. Additional details on the data including sources are found in Section A1 of the Appendix.

Across all specifications, we use benchmark-consistent official UST flow measures following Bertaut and Tryon [2007] and Bertaut and Judson [2014] and also the raw TIC flow data for robustness. For reference, the baseline results refer to those estimated using benchmark-consistent flows because of the well-known limitations of raw TIC data.[15]. These flow measures are plotted in Figure 3.

[13] Total public debt corresponds to FRED data series "GFDEBTN" and debt held by Federal Reserve banks corresponds to series "FDHBFRBN". These measures include all Treasury debt (bills, notes, bonds).

[14] We also considered realized bond yield volatility estimated as monthly standard deviations using daily changes in the 10-year treasury yield as an additional risk proxy, but its inclusion does not change the results whatsoever, and its corresponding regression coefficient is statistically insignificant.

[15] The TIC data cannot differentiate official flows when the transaction goes through a third-party intermediary and therefore the TIC reported flows can be considered a lower-bound estimate. Second, TIC data tend to overstate purchases of some securities (such as U.S. Agency bonds). For these reasons, we also consider the benchmark-consistent flows data constructed by combining the more accurate annual benchmark survey data with the TIC flows data on foreign holdings

We augment the specification in Warnock and Warnock [2009] by including L lags of foreign official UST purchases. The literature usually considers a 12-month rolling sum of foreign UST purchases as the main covariate of interest, with the coefficient interpreted as a 'long-run flow effect'. However, this is a special case of (3) where $L = 11$ and $\theta_l = \theta$ for all $l = 0, ...11$. Therefore (3) generalizes the previous modeling approaches by allowing for the possibility of different coefficients on each lag of ΔFO_t. The estimated cumulative or long-run impact of UST purchases or sales on U.S. yields over 12 months is given by $\sum_{l=1}^{11} \theta_l/(1-\phi_0)$. In the case where $\phi_0 = 0$ (excluding the lagged dependent variable), this is just the sum of coefficients across all L lags of ΔFO_t plus the contemporaneous flows. Further details on estimation and inference are in Section A5 of the Appendix.

3.2 OLS Estimation Results

Table 1 reports the estimation results. The first set of estimates do not control for common foreign yield factors. Broadly speaking the resulting estimates are consistent with the literature: higher short-term yields pass through as higher long-term yields, positive GDP and inflation forecasts are associated with higher 10-year yields, while higher risk premia (VIX) and structural budget surpluses are associated with lower 10-year yields. All regressions, estimated in levels, are tested for stationarity of the residuals via Augmented Dickey-Fuller (ADF) tests, which strongly reject the null of unit root in all cases, suggesting the presence of cointegration in the 10-year yield and supporting the application of a level specification.

The coefficient estimate reported for foreign official flows is for the sum of coefficients over the latest 12 months ($l = 0, ..., 11$), and it is significantly negative. Referring to the results using benchmark-consistent flows, a 12-month foreign official sale of USTs amounting to 1 percent of debt is associated with 10-year yields rising about 34.8 basis points. Assuming marketable U.S. debt of \$18 trillion as it read in 2017, a \$100 billion dollar sale of USTs by foreign officials over 12 months would be associated with a 19 basis point rise in 10-year yields. Estimates under raw TIC flows suggest an impact from a \$100 billion flow of roughly

Table 1: 10-Year U.S. Yield Regressions

	Dependent Variable: 10Y U.S. Yield							
	TIC Flows				Benchmark-Consistent Flows			
3M U.S. Yield	0.258***	(0.029)	0.319***	(0.023)	0.372***	(0.033)	0.388***	(0.027)
1Y GDP Forecast	0.017	(0.067)	0.040	(0.066)	0.488***	(0.100)	0.444***	(0.106)
10Y Inflation Forecast	0.106	(0.657)	-0.630	(0.431)	0.347	(0.608)	-0.139	(0.407)
1Y Inflation Forecast	0.047	(0.078)	0.117**	(0.068)	-0.057	(0.065)	-0.050	(0.060)
VIX	-0.021***	(0.004)	-0.015***	(0.005)	0.009	(0.006)	0.005	(0.006)
Budget Surplus	-0.082***	(0.037)	-0.015	(0.028)	-0.054	(0.039)	-0.013	(0.033)
12M Foreign Official Flows	-1.017***	(0.151)	-2.024***	(0.094)	-0.348*	(0.199)	-1.105***	(0.145)
10Y Global Yield			0.810***	(0.064)			0.689***	(0.065)
3M Global Yield			-0.391***	(0.060)			-0.248***	(0.060)
Adj. R^2	0.909		0.940		0.916		0.942	
T	266		266		240		240	
ADF Statistic	-4.798***		-4.071***		-5.282***		-4.465***	

Regression estimates from Equation 3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Coefficients and standard errors for foreign official flows are estimated over the sum of coefficients on ΔFO_{t-l} where $l = 0, ...11$ over the last 12 months. Foreign official flows variable is scaled by U.S. marketable debt lagged 12 months. Regressions include an intercept term and linear time trend. Augmented Dickey-Fuller (ADF) test on regression residuals reject the null hypothesis that residuals are nonstationary. The sample periods under TIC flows and benchmark-consistent flows are January 1999 to February 2021 and January 1999 to December 2018, respectively. Benchmark-consistent flows data are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

55 basis points. These estimates align closely with several reported in the literature despite different sample periods, methods and data (Table A.1).

Next we introduce the 10-year and 3-month common foreign yield factors as controls. The estimate on foreign official flows becomes significantly more negative, from -0.348 to -1.105 (using benchmark-consistent flows). This is consistent with an omitted variable bias hypothesis in the absence of controlling for common factors as we discussed above. Under \$18 trillion marketable debt, this implies that a \$100 billion sale of USTs by foreign officials is associated with 60.5 basis point rise in 10-year yields, which is three times larger than the estimated 19 basis point impact when only controlling for domestic factors. Most of this impact is through the term premium (Table A.7). After controlling for the common factors, the same 1 percentage point increase in foreign sales of USTs is associated with an increase of 76 basis points in the term premium (about 70% of the impact on the nominal 10-year yield, and compared to a term premium increase of 25 basis points in the absence of global factors).

Figure A.7 in Section A5.2 of the Appendix traces the historical contribution of foreign official UST flows (as a percentage of marketable debt) on the U.S. 10-year yield over the sample period (through February 2021 for TIC flows and through December 2018 for benchmark-consistent flows). The results confirm previous studies which document substantial yield compression attributed to foreign demand for USTs particularly in the early 2000s.

4 Identifying a UST Shock

The baseline regression, while useful to assess the scope for omitted variable bias, does not control for the likely endogeneity given the dynamics and feedbacks observed between international capital flows and interest rates. To capture these dynamics and address endogeneity, a recent literature uses VARs, but existing estimates vary widely.[16] Unlike previous studies, we control for foreign common factors in the VAR analysis and address identification in novel manner not yet considered by the literature. Specifically, to estimate the dynamic impact of a foreign official UST purchase or liquidation shock, we estimate a VAR with U.S. yields of differing maturities and (benchmark-consistent) foreign official UST flows as follows:

$$\mathbf{Y}_t = \boldsymbol{\beta}'\mathbf{Y}_{t-l} + \boldsymbol{\Gamma}'\mathbf{X}_t + \mathbf{u}_t, \tag{4}$$

where

$$\mathbf{Y}_t = [\Delta FO_t, y_{us,t}^{3M-FF}, y_{us,t}^{2Y-FF}, y_{us,t}^{5Y-FF}, y_{us,t}^{10Y-FF}, y_{us,t}^{30Y-FF}],$$

$$\mathbf{X}_t = [\mathcal{Y}_{g,t}^{3M}, \mathcal{Y}_{g,t}^{10Y}, \Delta GDP_t^{E[t+1]}, \pi_t^{E[t+1]}, \pi_t^{E[t+10]}, VIX_t, surplus_t, \mathbf{D}_t].$$

[16] Ayanou [2016] estimates statistically insignificant effects from the VAR, Wolcott [2020] estimates a 17 basis point impact, Fang and Liu [2019] estimates a 50.5 basis point impact.

The VAR represented in (4) can be viewed as the dynamic extension of the static regression model in (3). It is parsimonious in its assumptions and transparent.[17] We consider 3-month, 2-year, 5-year, 10-year, and 30-year U.S. yields. These yields enter as spreads relative to the Federal Funds rate denoted FF which effectively de-trends these level variables while accounting for the stance of Fed monetary policy. The matrix $\mathbf{D}_t$ includes a set of three dummy variables that remove major residual outliers that produce wide error bands in the bootstrap procedure used to construct the confidence set around the IRF estimates. The three dummies are February 2008, September 2008 and October 2008 coinciding with the collapse of AIG and Bear Stearns and the Lehman Brothers crash.[18] The VAR is estimated with 4 lags similar to Bernanke et al. [2004], but the results are robust to alternative lag length assumptions.

4.1 Identification Through Heteroskedasticity

We address the endogeneity problem by identifying structural shocks in the VAR above by exploiting the shift in the time series' variance after the Lehman crash in September 2008. This *identification through heteroskedasticity* approach was initially proposed in Rigobon [2003] and recently applied by Brunnermeier et al. [2021] in a VAR setting. The basic idea is as follows. Suppose the covariance matrix of the residuals $\mathbf{u}_t$ differ before and after September 2008:

$$E(\mathbf{u}_t\mathbf{u}_t') = \begin{cases} \Sigma_1, \text{ for } t = 1, ..., t_{Sep2008} - 1 \\ \Sigma_2, \text{ for } t = t_{Sep2008}, ..., T, \end{cases} \tag{5}$$

so that $\Sigma_1 \neq \Sigma_2$. The two covariance matrices can be expressed as $\Sigma_1 = BB'$ and $\Sigma_2 = B\Lambda B'$ where Λ is a diagonal matrix with strictly positive elements λ_{ii}, $i = 1, ..., 6$, and the

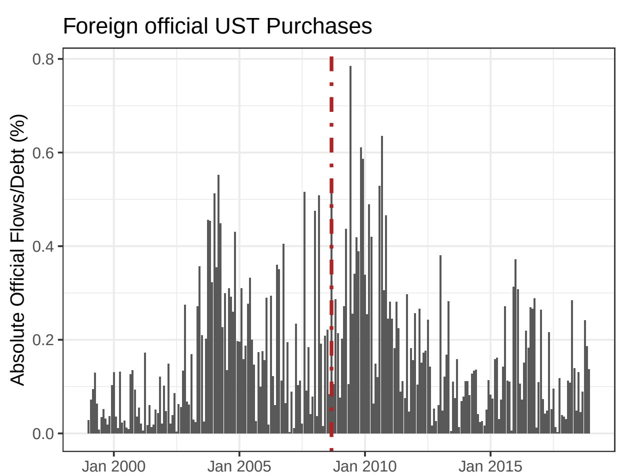
[17] Section A5 discusses the specification of this VAR in more detail and compares it against alternative term structure models used in the literature.

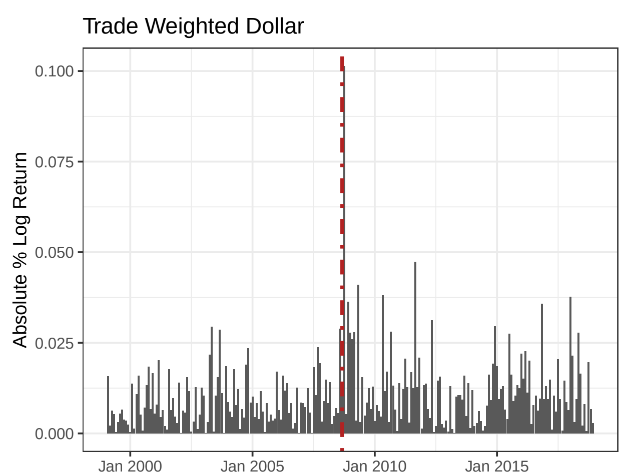
[18] We also considered including dummies for the September 2001 terrorist attacks and 2013 'Taper Tantrum'. Including these additional indicators tightens the confidence bands further.

relationship between the unconditional residual covariance matrix and that of the structural shocks (ζ_t) is $\Sigma_{\mathbf{u}} = B\Sigma_\zeta B'$. The first regime variances are normalized to equal one and the matrix Λ characterizes the relative change in variances in second regime. The structural shocks are identified if all elements of Λ are distinct.

A structural break in the variance of the time series' is crucial for the validity of identification through heteroskedasticity. Typically these breaks are based on external knowledge of material historical events. In our case, we choose a break in September 2008 marking the Lehman crash because of the well-documented shift in the level volatility of international capital flows before and after the global financial crisis. López and Stracca [2021] documents that capital flows diminished abruptly since the 2008 crisis and that the composition of flows shifted from bank flows to investor portfolio flows. Forbes and Warnock [2021] show that factors driving capital flows shifted from risk appetite to commodity prices, and extreme capital flow movements decreased since the crisis. Erik et al. [2020] find that the role of the U.S. Dollar has gained importance since the 2008 crisis as a risk factor. Consistent with these views, we find that the volatility of foreign official UST flows changed since 2008.

Figure 4: Absolute Foreign Official UST Flows and Absolute U.S. Trade-Weighted Dollar Returns





Dashed vertical line is September 2008. Pre/post September 2008 variance ratio of foreign official UST purchases and U.S. Dollar returns are 1.81 and 2.64, respectively. Both are statistically significant at the 1% level.

Figure 4 left panel shows that the magnitude of foreign official UST flows changed significantly in the post-2008 period. The monthly standard deviation of foreign official flows increased from 0.17% to 0.23%. F-tests for the ratio of pre/post September 2008 variances indicate that the shift is statistically significant at the 1% level. The right panel plots monthly logged U.S. traded-weighted Dollar returns. On an annualized basis, the volatility of the U.S. Dollar increased from 3.77% to 6.14% in the period after September 2008 and again the change in variance is significant at the 1% level.[19]

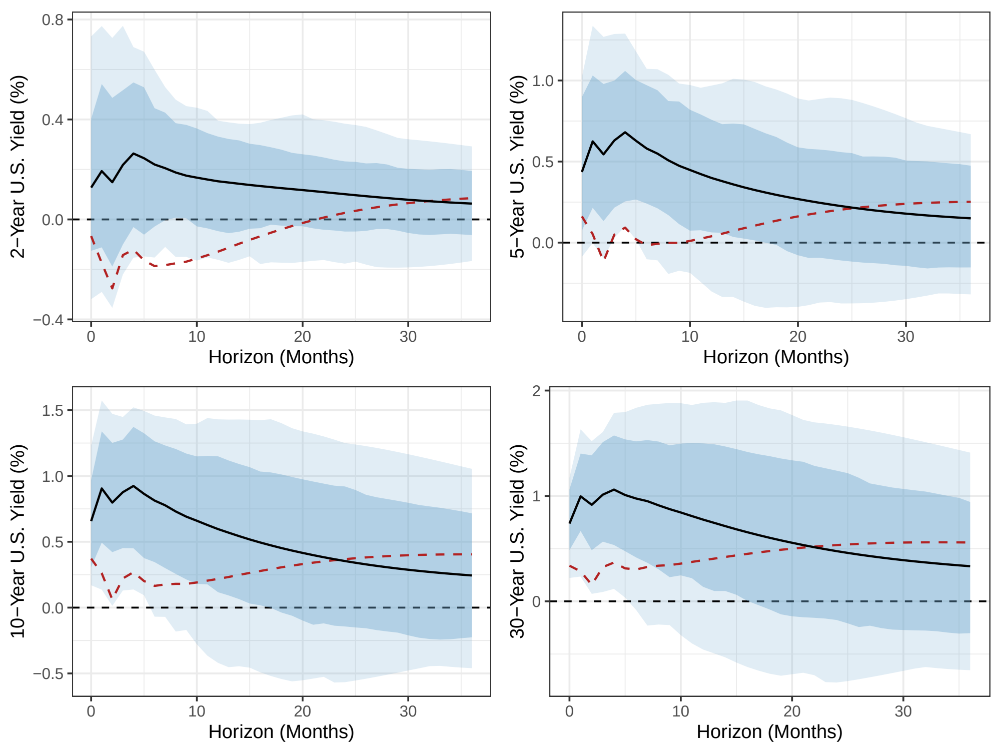
4.2 Estimated Impulse Response Functions

Figure A.8 traces the impulse response functions (IRF) from an official UST sale the size of 1 percent of U.S. marketable debt in one month on 2-year, 5-year, 10-year, and 30-year nominal Treasury yield spreads over the Federal Funds rate (the IRF from an official UST purchase is simply the equivalent IRF but flipped). Note that under identification through heteroskedasticity, the shape of the IRFs are identical across regimes, but the size of the effects can differ. Instead of choosing a specific regime, Figure A.8 traces IRFs scaled by the *unconditional* variance of the reduced form residuals over the entire sample period. The U.S. yield response increases with maturity, consistent with the existing evidence suggesting that shorter maturity yields are largely determined by domestic U.S. monetary policy and an effect that transmits through the term premium. The 3-month yield response is insignificant and not shown to conserve space. The dashed red line reflects an equivalent IRF in a model that does not incorporate contemporaneous foreign common factors and only conditions on domestic factors, and is shown to be substantially weaker (and often insignificantly different from zero).

Following a foreign official UST sale of 1% of marketable Treasury debt, 2-year yields rise about 16 basis points on impact, while 5-year yields tend to rise about 50 basis points, and

[19]These significant regime changes in volatilities persist under different break dates (e.g. if we exclude the crisis period altogether) and we find similar changes in UST flow volatility when inspecting large UST holding countries such as China.

Figure 5: Impulse Response of U.S. Yields over the Federal Funds Rate to a Foreign Official UST Sale (1% of Debt)



Impulse responses from a VAR(4) specified as in Equation 4 using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. Structural shocks are identified by heteroskedasticity. Light and dark shaded regions refer to 90% and 68% bootstrapped confidence intervals based on 1,000 bootstrapped samples. Dashed line reflects impulse response from VAR(4) that excludes global yield factors from the set of contemporaneous controls. Unit variance shocks are scaled using unconditional variance of the reduced form residuals to arrive at the impact from a 1% sale of USTs.

10 and 30-year yields increase about 65 and 75 basis points, respectively. These effects peak at about 5 months following the sale by foreign officials before slowly reverting back to the steady-state.[20] As a result, foreign official UST demand affects not just the level, but also the slope of the yield curve.[21] While the estimated effects are large relative to the literature, they are consistent with biases (global confounding factors and simultaneity) which cause prevailing estimates to be understated when not properly accounted for. The magnitude of the estimated effect is discussed further in Section 5 below, where we compares estimates of the impact of foreign official demand for USTs with estimates found in the Quantitative Easing literature (i.e. domestic official demand for USTs).

5 Controlling for Federal Reserve purchases of U.S. Treasuries

Although difficult to measure, Federal Reserve forward guidance and large scale asset purchase (LSAP) policies can play an important role in shaping longer-term U.S. yields via both direct and information channels. Since these unconventional policies are typically administered when conventional policy rates are constrained by the zero lower bound, it is also important to condition on these variables to account for potential changes in yield dynamics arising from shifting monetary regimes. Recent advances in Swanson [2021] extend upon high-frequency identification literature of monetary policy shocks to better extract unconventional policy shocks specifically by examining high-frequency movements across financial asset classes around FOMC announcements.

[20] The impact of a \$100 billion UST sale (or purchase) of USTs over one month would then be roughly 35.75 and 45.75 basis points on the 10-year and 30-year yield, and 27.5 basis points on the 5-year yield. If common foreign factors were omitted from the VAR, the impact falls substantially. On the 10-year yield, the impact falls to about 16 basis points.

[21] While consistent with Kaminska and Zinna [2020], the larger impact on longer-term maturities is inconsistent with the fact that foreign official UST portfolio has an average duration of about 4 years. Reconciling this puzzle would likely require more granular capital flow data, disaggregated by duration or at least maturity.

We leverage the forward guidance and LSAP shocks of Swanson [2021] to control for any further forward-looking domestic forces shaping U.S. yields which the standard controls might not capture. Moreover, conditioning on unconventional policy helps to account the shift in monetary regimes since the 2008 Financial Crisis. Because these shocks are measured at the daily frequency, we aggregate them to the monthly frequency as a sum of the daily shocks within each month. First, the regression from (3) is augmented with two additional covariates reflecting forward guidance policy: the monthly forward guidance shocks themselves, and the cumulative sum of the monthly forward guidance shocks. We include the cumulative sum of shocks as an additional covariate to account for the fact that the dependent variable is the level of the 10-year U.S. yield rather than yield changes. Second, we augment the baseline regression with four additional variables to condition on both forward guidance and LSAP policies, the two shock series and the cumulative sums of these shock series. Table A.12 reports the impact of 12-month flows on U.S. 10-year yields after conditioning on forward guidance measures, finding that the estimates are nearly unchanged (strengthening very slightly).

After conditioning on both forward guidance and LSAP shocks, the estimates strengthen. The stronger effect size suggests that forward guidance and LSAP shocks are state-dependent and negatively correlated with foreign official UST demand. For instance, LSAP policies are typically administered during crises to bring down long-term yields, occurring during periods of economic weakness precisely with foreign officials are liquidating USTs. This implies that the presence of Federal Reserve interventions during cyclical downturns helps dampen the impact of foreign official UST selling on the U.S. yield, as Fed buying offsets foreign official selling during these downturns.

5.1 Comparing against estimates from the Quantitative Easing literature

A closely related literature studies the effects of Federal Reserve asset purchase programs to evaluate the effect of quantitative easing (QE) on U.S. interest rates. Estimates across these studies vary, but are generally smaller than the estimates of the effect of foreign official purchases. Broadly speaking, a $100 billion bond purchase via QE has been linked to compression in long-term U.S. yields ranging from 5 to 15 basis points [Gagnon et al., 2011; D'Amico and King, 2013; Bonis et al., 2017; Swanson, 2021]. These estimates are substantially smaller than the 37.75 basis point effect we report from the VAR estimates here in Section 4 but are in line with the VAR estimate of 16 basis points when global factors are omitted.

There are several reasons why domestic and foreign official purchases of USTs might affect U.S. bond yields differentially. One important factor is the endogeneity of QE programs. The Federal Reserve tends to announce bond purchases when bond market liquidity is quickly disappearing. Similar to most studies on foreign official demand and U.S. yields, most studies on the impact of domestic official asset purchases also do not account for the role of global cyclical factors which almost certainly influence market conditions. As such, Fed purchases may be meeting large selling pressure, which would lead to a smaller price impact on U.S. yields compared to the same bond purchase amid normal market conditions. The main results from this study are consistent with this 'liquidity view' but in the context of foreign official UST purchases.

Second, the economic implications of foreign and domestic official asset purchases can be starkly different. While domestic asset purchases serve essentially as an asset exchange (Dollar for Dollar) where primary changes are compositional (e.g. lower aggregate duration), foreign flows entering the U.S. economy provide additional resources and purchasing power [Kohn, 2016]. Beltran et al. [2013] mentions other important reasons why U.S. yields may be more sensitive to foreign official demand than domestic official demand. These include the expectation that QE policies are relatively temporary, or expectations of subsequent QEs becoming priced in bond markets even prior to their announcement. Finally unlike foreign official UST purchases, domestic QE programs tend to be associated with greater inflation uncertainty, which would put upward pressure on yields and offset some of the downward pressure brought on by Federal Reserve asset purchases.

6 Robustness Checks

In this section we discuss a variety of robustness checks on the baseline regression specification detailed in Section A5 of the Appendix. We then explore alternative identification strategies for the VAR specification, with additional detail in Sections A6 and A7 of the Appendix.

6.1 Alternative regression specifications

For robustness, the large changes in estimated effect sizes upon introducing foreign yield factors as controls are similarly found in specifications when assigning the term premium as the dependent variable (Table A.7), that include a lagged dependent variable (Table A.8), and when scaling foreign official flows by U.S. GDP instead of marketable debt (Table A.9) as done in Warnock and Warnock [2009]. In the case of including a lagged dependent variable and scaling by GDP, the absolute size of the effect is larger than those reported in the baseline results which do not include a lagged dependent variable and scale flows by marketable debt. Table A.10 extends the sample to start from 1990 and finds that the results are generally robust to the lengthening the sample period though diminishing slightly.

We also consider regression specifications that replace 12-month UST flow with the level of foreign official UST holdings, to estimate the permanent or long-run 'stock effect' (Table A.11). Consistent with Beltran et al. [2013], the permanent effect of a change in foreign official UST holdings is quantitatively smaller than the 12-month flow effect. A one percentage point decrease in the stock of debt held by foreign officials is associated with 10-year U.S.

yields rising by 10.5 basis points after controlling for foreign yield factors, which is almost double the effect size when ignoring global factors (5.5 basis points). The estimates are also robust to constructing the foreign yield factors based on debt weights, equal weights, or PCA weights (Table A.12 in the Appendix). Section A5.3 of the Appendix reports long-run estimates from an error-correction model on monthly 10-year yield changes, and effect sizes are roughly two to three times larger after accounting for foreign yield factors.

6.1.1 Two-way feedback effects between U.S. and global yields

In the baseline regression specification, the foreign yield factors, $\mathcal{Y}_t^{3M}$ and $\mathcal{Y}_t^{10Y}$ are assumed to affect U.S. yields but there is sufficient evidence to believe that U.S. yields also affect foreign yields. While there is no simple solution to separate these two-way feedback effects from U.S. yields to foreign yields and vice versa, we consider several approaches which try to remediate this problem. First, foreign yield factors enter the regression lagged by one month ($\mathcal{Y}_{g,t-1}^{3M}$, $\mathcal{Y}_{g,t-1}^{10Y}$) instead of contemporaneously. Second, we consider modified foreign 10-year and 3-year yield factors which are uncorrelated with the U.S. 10-year and 3-year yield, respectively. These orthogonalized factors are recovered by regressing the foreign 10-year and 3-year yield factors on the 10-year and 3-year U.S. yield:

$$\mathcal{Y}_{g,t}^{10Y} = \alpha + \beta y_{us,t}^{10Y} + \mathcal{E}_{g,t}^{10Y}, \qquad \mathcal{Y}_{g,t}^{3M} = \alpha + \beta y_{us,t}^{3M} + \mathcal{E}_{g,t}^{3M}, \tag{6}$$

where $\mathcal{E}_{g,t}^{10Y}$ and $\mathcal{E}_{g,t}^{3M}$ are the regression residuals, and the component of the foreign yield factor which is uncorrelated with the U.S. 10-year yield. We then use these residuals in place of $\mathcal{Y}_{g,t}^{10Y}$ and $\mathcal{Y}_{g,t}^{3M}$ in the baseline regression specification (3). This approach can be viewed as relatively conservative since by design it purges all contemporaneous correlation between the U.S. and foreign yields which may arise from either 1) U.S. yields affecting foreign yields, 2) the joint response of U.S. and foreign yields to some third factor, 3) correlation arising from foreign yields affecting the U.S. yield, thereby 'over-purging' in a sense. However, it does not necessarily remove the correlation between foreign yields and foreign official flows, which is another important source of correlation we wish to control for. Finally, we also consider the 10-year U.S.-foreign yield differential as a dependent variable, $y^{10Y}_{us,t} - \mathcal{Y}^{10Y}_{g,t}$ which is an implicit restriction on the baseline regression (3) with $\phi_3 = 1$. This specification aims to explain the 10-year U.S.-foreign yield spread rather than 10-year nominal yields alone.

Table A.12 in the Appendix reports the long-run estimates from using lagged foreign factors ($\mathcal{Y}^{3M}_{g,t-1}$ and $\mathcal{Y}^{10Y}_{g,t-1}$), the orthogonal foreign yield factors ($\mathcal{E}^{3M}_{g,t}$ and $\mathcal{E}^{10Y}_{g,t}$), or when replacing the dependent variable with the U.S.-foreign 10-year spread, respectively. Under all three specifications, the long-run impact of foreign official flows on U.S. yields remains significant and large. Using the benchmark-consistent foreign official flow measure, a one percentage point drop in flows relative to marketable debt is associated with a 103.7 basis point rise under the lagged specification (*"GDP-weighted, lagged"*), a 47.7 basis point rise under the orthogonal specification (*"GDP-weighted, orthogonal"*), and a 144.8 basis point widening of the U.S.-foreign 10-year spread (*"GDP-weighted, $y^{10Y}_{us,t} - \mathcal{Y}^{10Y}_{g,t}$*). This translates to a 57.2, 26.2, and 79.64 basis point impact per $100 billion foreign official sale over the long-run, respectively.

6.1.2 CIP deviations, Geopolitical risk, World economic activity, Emerging market shocks

Because the foreign yield factors are recovered from local currency bond yields, they do not directly account for the U.S. Dollar cross-currency basis, i.e. the Treasury premium. The U.S. Treasury premium is thought to reflect the special safety and liquidity feature of physical U.S. Treasury bonds over synthetic replicates, and this premium may be an important source of variation in U.S. yields that may be potentially confounded with foreign official demand for USTs. To control directly for the basis, we use the G10 government bond yield CIP deviations of Du and Schreger [2016] and Du et al. [2018] to construct average 3-month and 10-year U.S. Treasury premia, which are included as additional controls in the baseline regression (Table A.12, *"Controlling for CIP Deviations"*). The estimated long-run impact of foreign official UST flows on U.S. 10-year yields falls slightly but remains large and significant. Referring to the benchmark-consistent flow measures, a fall in flows relative to marketable debt of 1 percent is associated with a rise in U.S. yields of 82.6 basis points (45.4 basis points per $100 billion flow). Additional details are provided in the Appendix in Section A5.

Political risk and global uncertainty can also jointly affect the incentive to accumulate international reserves by foreign institutions, and the pricing of U.S. yields. To account for the role of geopolitical risk, the news-based index of Caldara and Iacoviello [2018] is used, which counts the frequency of words related to geopolitical tensions in 11 leading international newspapers. Geopolitical risk enters significantly as a conditioning variable, and Table A.12 shows that the impact of foreign official UST flows on yields decreases slightly, though remains large and significant.

In addition to the foreign yield factor, we add a concrete measure of non-U.S. global economic activity taken from the Dallas Fed Database of Global Economic Indicators. Specifically, as an additional factor we include year-on-year world (ex. U.S.) industrial production (IP) growth, which is sampled at the monthly frequency from 40 countries, both advanced and emerging economies. Table A.12 shows that the impact of foreign official UST flows on U.S. 10-year yields remains close to the baseline estimate after conditioning additionally on world economic activity.

In the uncommon case that emerging market shocks spill over to the United States, it would be important that foreign factors also capture changes in emerging market conditions which induce central banks of these countries to buy or sell USTs. While advanced economy yield factors should account for a reasonably large portion of emerging market financial conditions, the measure may be incomplete. As robustness check, emerging market yields are specifically controlled for alongside the advanced economy yield factors using monthly changes in the EMBI sovereign spread index. Results for long-run estimates are shown in Table A.12, which remain close to the baseline results.

6.2 Alternative VAR Identification Strategies

In this section, we present alternative estimates of the price impact of foreign official demand under alternative identification strategies.

6.2.1 Short-run exclusion restrictions

A common approach toward identifying structural shocks from a VAR is through a Choleky decomposition. Under this scheme, the structural shocks to foreign official flows are identified recursively from the VAR reduced-form residuals $\mathbf{u}_t$ using short-run zero restrictions. Specifically, foreign official flows are ordered first in the VAR, affecting U.S. yields contemporaneously, but U.S. yields can only affect official flows with a one-month lag. This restriction is conceptually consistent with contemporaneously inelastic demand of foreign officials (whose demand for USTs does not instantaneously respond to U.S yields). Figure A.8 of Section A5 in the Appendix plots these impulse response functions which are both qualitativey and quantitatively consistent with the baseline results. However it is worth noting that the intial impact under this identification scheme is insignificant, and the impact on U.S. yields materializes with a delay.

To further guard against simultaneity between foreign official flows and U.S. yields and global confounding factors, we also consider recursive ordering where foreign official flows are ordered last. Figure A.9 of Section A5 in the Appendix shows that the IRFs from (4) are robust and of similar size when ordering foreign official flows last in the structural VAR instead of first. By ordering foreign official flows last, they no longer have a contemporaneous impact on U.S. yields.

Finally, figure A.10 of the Appendix plots IRFs following a sale of USTs by foreign *private* investors form a VAR which simply augments (4) with foreign private UST flows scaled by marketable debt. Because foreign private flows are likely not inelastic the way foreign official flows are, foreign private flows are ordered last in the Cholesky decomposition to guard against simultaneity with U.S. yields. Moreover, greater demand elasticity implies smaller price impact, so if foreign official flows are relatively inelastic, the estimated price impact of foreign private flows should be smaller. A sale of USTs the size of 1 percentage of debt by foreign private investors induces a rise in yields of about 12 basis points in the 5-year Treasury and 14 basis points in the 10-year Treasury. Consistent with relatively elastic foreign private demand (and inelastic official demand), these effect sizes are roughly half those from foreign official UST flows. This result suggests that global confounding factors may be a more important source of endogeneity than simultaneity when studying the effect of foreign official UST flows on U.S. yields.

6.2.2 Demand from China

In addition, Section A7 in the Appendix takes a novel two-pronged identification strategy in the spirit of external instrument approaches toward identification of SVARs. This approach exploits *lagged* variation in China's exchange rate volatility while continuing to condition on foreign yield factors in the VAR. The motivating idea being that China's transitioning exchange rate policies over the last two decades determined to a large extent their official demand for USTs. The baseline results hold under this alternative empirical strategy and suggest that U.S. yields are sensitive to official demand linked to China. A back-of-the-envelope calculation implies a 20 to 50 basis point impact per $100 billion in Chinese official UST flows.

7 Concluding Remarks

Both OLS estimates and VAR impulse responses identified through heteroskedasticity demonstrate that the impact of foreign official demand for USTs on U.S. yields is substantially understated (by 30-50%) when foreign common factors in yields are omitted and only domestic factors are controlled for. A key reason why previous estimates of the impact of foreign official demand for USTs on U.S. yields may be biased is because they do not account for the cyclical nature of U.S. yields and foreign demand for USTs. During global economic expansions, downward pressure on U.S. yields exerted by increased official demand for USTs from abroad is met with upward pressure exerted by improving U.S. and global economic fundamentals and weaker UST demand from domestic private investors. We also find that the impact of foreign official demand for USTs on U.S. yields increases further after accounting for Federal Reserve LSAPs, which tend to occur during crises when foreign officials are selling USTs. These results bear three important policy implications. First, the impact of UST liquidations or purchases of the black-swan type – those not driven by cyclical forces – may have a substantially larger impact on U.S. Treasury market functioning than previously thought. Second, pro-cyclical foreign official demand may be an important source of Treasury market liquidity. Third, the effectiveness or price impact of government asset purchase programs is likely to depend on market conditions.

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Appendix

The Appendix is organized as follows: Section A1 provides detail on relevant data and sources. Section A2 presents evidence of a global factor in bond yields, and shows that global yield factors are not spanned by traditional U.S. level, slope, and curvature term structure factors when explaining U.S. 10-year and 3-month yields. Section A3 compares global yield factors constructed under alternative weight schemes. Section A4 presents supporting evidence suggesting that global yield factors reflect global cyclical conditions, which jointly influence U.S. yields, foreign official demand for USTs, and U.S. private investor demand for USTs. Section A5 provides robustness checks for the baseline regression specification and Section A6 and A7 discuss the VAR model and reports results under alternative identification strategies including Cholesky decomposition and exploiting foreign UST demand from China.

A1 Data

A summary list of estimated impacts from a \$100 billion purchase of USTs by foreign officials is found in Table A.1.

All data used in this analysis is publicly available and collected from a variety of sources. First, to construct global bond yield measures, monthly long-term and short-term government bond yields across countries are from the Organization for Economic Co-operation and Development (OECD) database. Long-term and short-term yields correspond to 10-year and 3-month maturity bonds, respectively. Along with U.S. interest rates, interest rate data are collected for 19 countries: Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, The United Kingdom, Ireland, Italy, Japan, Luxembourg, Netherlands, Norway, New Zealand, and Sweden. Quarterly U.S. dollar values of nominal

Table A.1: The Impact (in Basis Points) of a $100 Billion Foreign Purchase or Sale* of U.S. Treasury Securities on the U.S. Long-Term Treasury Yield: Estimates from Previous Studies

Study	**Impact**	**Measurement**	**Sample Period**
Bernanke et al. (2004)	-66	Japanese off. intervention (daily)	1/3/2000-3/3/2004
McCauley and Jiang (2004)	-70 to -100	For. off. flows of Treasury and Agency securities (weekly)	2000M01-2004M01
Rudebusch et al. (2006)	No effect	12-month for. off. flows into Treasury securities (% debt)	1990M05-2005M12
Bandholz et al. (2009)	-12	For. total holdings of Treasury securities (% debt)	1986M01-2006M06
Craine and Martin (2009)	-61	For. off. holdings of Treasury securities (% personal income)	1/1/1990-12/31/2003
Warnock and Warnock (2009)	-34 to -68	12-month for. off. flows into Treasury and Agency securities (% of GDP)	1984M01-2005M05
Bertaut et al. (2012)	-13	For. off. holdings of Treasury and Agency securities (% debt)	1980Q1-2007Q2
Kitchen and Chinn (2012)	-55	For. off. holdings of Treasury notes and bonds (% of GDP)	1979-2010
Beltran et al. (2013; long-run,flow)	-39 to -62	12-month for. off. flows into Treasury securities (% of debt)	1994M01-2007M06
Beltran et al. (2013; short-run, flow)	-46 to -50	For. off. flows into Treasury securities (% of debt)	1994M01-2007M06
Beltran et al. (2013; long-run, stock)	-17 to -20	For. off. holdings of Treasury securities (% of debt)	1994M01-2007M06
Martin (2014a)	-76	For. off. flows of Treasury notes and bonds (% of debt)	1985M01-2011M06
Martin (2014b)	-100	Chinese exchange rate policy announcements (daily)	1/1/2005-4/30/2014
Ayanou (2016)	-50	For. off. holdings of Treasury notes and bonds (% of GDP)	1982M01-2009M03
Gerlach-Kristen et al. (2016)	-100	Japanese off. intervention (daily)	1/15/2003-3/17/2004
Kohn (2016)	-50	For. off. holdings of Treasury notes and bonds (% of GDP)	2000Q1-2007Q4
Csonto and Tovar (2017)	-1.8	For. off. holdings of Treasury notes and bonds (% of debt)	2003M01-2015M12
Fang and Liu (2019)	-50.5	For. off. holdings of Treasury notes and bonds (% of debt)	1985M01-2007M06
Kaminska and Zinna (2020)	-4	For. off. holdings of Treasury notes and bonds (% of debt)	2001M01-2012M11
Wolcott (2020)	-17	For. off. flows of Treasury notes and bonds (% of debt)	1985M01-2014M08
Zhang and Martinez-Garcia (2020)	-5 to -11	For. off. holdings of Treasury notes and bonds (% of debt)	1986M01-2014M12
This study: long-run, flow	-60.5	12-month for. off. flows of Treasury notes and bonds (% of debt)	1999M01-2018M12
This study: short-run, flow	-27.5	For. off. flows of Treasury notes and bonds (% of debt)	1999M01-2018M12
This study: long-run, stock	-5.8	For. off. holdings of Treasury notes and bonds (% of debt)	1999M01-2018M12

*Effect of UST sale would be of the same magnitude but opposite sign. The value of the scaling variable (GDP, debt, etc.) chosen to recover the estimated impact is based on the values of the scaling variable within the sample period of study. Because these scaling factors tend to rise over time, the estimated impact of a $100B purchase is more likely to be smaller in more recent periods. Estimates across studies can differ due to differences in scaling variable, measures of foreign purchases of Treasuries, sample period, modeling approach (short-run versus long-run impact), dependent variable (10-year yield, 5-year yield, real or nominal yield, term premia, mortgage rates). Abbeviation 'for. off.' refers to 'foreign official'. 'No effect' means the estimated impact from the study was statistically indifferent from zero. '*This study*' estimates based on benchmark-consistent official flows data of Bertaut and Tryon [2007] and Bertaut and Judson [2014]. See [Hoelscher, 1986; Bernanke et al., 2004; McCauley and Jiang, 2004; Rudebusch et al., 2006; Bandholz et al., 2009; Craine and Martin, 2009; Warnock and Warnock, 2009; Bertaut et al., 2012; Kitchen and Chinn, 2011; Beltran et al., 2013; Martin, 2014a,b; Sierra, 2014; Ayanou, 2016; Gerlach-Kristen et al., 2016; Kohn, 2016; Csonto and Tovar, 2017; Fang and Liu, 2019; Kaminska and Zinna, 2020; Wolcott, 2020; Zhang and Martínez-García, 2020]. Extensions to other countries beyond the U.S. include Carvalho and Fidora [2015], Blattner and Joyce [2016], Arslanalp and Poghosyan [2016], Inoguchi [2021].

gross domestic product (GDP) is also taken form the OECD for these countries, and these quarterly observations are converted to monthly frequency by imputing observations between fiscal quarter months (January, April, July, October) using the previous most recent value. Quarterly U.S. dollar values of national public debt are taken from the Bank for International Settlements (BIS), and are similarly interpolated to monthly frequencies. The cross-country yield and GDP data are used in constructing GDP-weighted global yield factors (and as a robustness check, yield factors are also constructed using national debt data). Long-term yield data for Luxembourg are missing from June 2007 to April 2010. We impute values for these missing values with fitted yield estimates from a regression using the observed data of Luxembourg bond yields on the yields of several European countries: Germany, the U.K., France, Spain, Belgium, Finland, Ireland, Italy, Netherlands, and Austria. Short-term yields for Japan are missing from January 1999 to March 2002. For these dates with missing values, we use the rates paid on Japanese 3-month certificates of deposits which are available on the Federal Reserve Economic Database (FRED). Monthly U.S. yields across several maturities (3-month, 6-month, 1-year, 2-year, 3-,year, 5-year, 7-year, 10-year, 20-year, 30-year) are taken from FRED.

Quarterly data on U.S. Treasury securities transactions for U.S. households (including hedge funds) and mutual funds comes from the Federal Reserve Financial Accounts. This data is used to show that private investor demand for USTs is counter-cyclical, unlike foreign official demand for USTs which is pro-cyclical.

To consistently estimate the impact of foreign official UST flows on U.S. interest rates, we need to control for determinants of U.S. bond yields beyond foreign demand. For this, we collect data on various controls used in the literature. Data for the daily VIX index along with U.S. quarterly nominal GDP, public debt outstanding, and public debt held by Federal Reserve banks (the latter three are used as scaling variables) are from FRED. Daily VIX readings are converted to monthly frequency by sampling the last value of each month. Daily U.S. yield data is also used (from FRED) to construct a monthly readings of realized

U.S. interest rate volatility as a secondary proxy for risk premia. Quarterly 1-year ahead GDP growth forecasts, 1-year ahead inflation forecasts and 10-year ahead average inflation forecasts are all taken from the Philadelphia Fed Survey of Professional Forecasters mean responses. 1-year ahead GDP growth forecasts for each period are computed as the average of 1, 2, 3, and 4-quarter ahead annualized forecasts. Structural budget surplus/deficit data is from the Congressional Budget Office (CBO). Quarterly data are linearly interpolated to monthly frequency. Data on daily Chinese Renminbi exchange rates vis-a-vis the U.S. Dollar are taken from FRED. To construct Figure 1, data on Treasury notes and bonds held by the public are taken from the U.S. Treasury Monthly Statement of the Public Debt (MSPD).

Table A.2: Summary Statistics

Statistic	T	Mean	St. Dev.	Min	Pctl(25)	Pctl(75)	Max
Long-term U.S. yield	266	3.400	1.394	0.620	2.270	4.515	6.660
Short-term U.S. yield	266	2.057	1.998	0.110	0.290	3.188	6.730
Long-term global yield factor	266	2.651	1.406	0.008	1.103	3.756	4.892
Short-term global yield factor	266	1.586	1.409	−0.287	0.230	2.722	4.272
1-year ahead U.S. GDP growth forecast	266	2.847	0.699	0.812	2.466	3.208	6.508
1-year U.S. inflation forecast	266	2.159	0.812	0.260	1.685	2.665	4.260
10-year average U.S. inflation forecast	266	2.395	0.145	2.070	2.260	2.510	2.700
Structural budget surplus	266	−3.195	2.073	−6.700	−5.092	−1.933	1.000
VIX index	266	20.287	8.183	9.510	14.023	24.490	59.890
Foreign official flows	266	−0.569	18.897	−61.203	−7.870	11.207	53.050
Foreign official flows (benchmark-consistent)	240	5.565	21.858	−57.161	−5.973	18.343	73.496

Data covers January 1999 to February 2021. Yield, GDP growth, inflation and VIX statistics are in percentages. Budget surplus statistic is as a percentage of GDP. Foreign official flows into U.S. Treasury notes and bonds are in billions of U.S. Dollars. Benchmark-consistent flows data based on Bertaut and Tryon [2007] and Bertaut and Judson [2014] and are available through December 2018. Quarterly data (forecasts and budget surplus) linearly interpolated to monthly frequency.

Finally, data on monthly purchases of U.S. Treasury notes and bonds by foreign officials come from the Treasury International Capital (TIC) system. The raw data is known to have some less than desirable features. The TIC data cannot differentiate official flows when the transaction goes through a third-party intermediary and therefore the TIC reported flows can be considered a lower-bound estimate. Second, TIC data tend to overstate purchases of some securities (such as U.S. Agency bonds). For these reasons, we also consider both the standard TIC flows data and the benchmark-consistent flows data constructed by combining the more accurate annual benchmark survey data with the TIC flows data on foreign holdings as proposed in Bertaut and Tryon [2007] and Bertaut and Judson [2014]. The benchmark-consistent flow data are available up to December 2018 while the TIC raw data sample ends in February 2021. Summary statistics for the data are reported in Table A.2.

A2 Common Foreign Yield Factors and Factor Strength

We consider monthly 3-month and 10-year government bond yields from the U.S. and 19 additional countries: Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, The United Kingdom, Ireland, Italy, Japan, Luxembourg, Netherlands, Norway, New Zealand, and Sweden from January 1999 to February 2021. Principal Component Analysis (PCA) reveals strong evidence of factor structure across global bond markets. Consider the long-term (10Y) bond yields. The first principal component (PC) of bond yield levels captures 91% of the variation across international bond markets and the first PC of bond yield changes captures 66% of the cross-country variation. Meanwhile, the second PC explains 6% and 9% in levels and changes, respectively. While the variation explained by the second PC is not trivial, the disproportionally large proportion of variance captured by the first PC motivates estimating at least a single global factor for long-term and short-term bond yields, respectively. Similarly, the average of the 190 pair-wise correlations between the 20 countries is equal to 0.90 for bond yield levels and 0.62 for bond yield changes, and these averages are highly statistically significant (t-statistics of 118 and 42, respectively for long-term yield level and changes). More formally, CD tests of Pesaran [2021] for cross-sectional dependence in bond yield levels and changes all strongly reject the null hypothesis of no cross-sectional dependence. These results are reported in Table A.3, and evidence of factor structure is similarly present in short-term yields.

We also can formally test factor strength proposed by Bailey et al. [2020] through determining the proportion of country long-term yields for which the global long-term yield factor enters significantly in a factor regression, while adjusting for multiple hypothesis testing. The

Table A.3: Evidence of Factor Structure in Global Bond Yields

	10Y Yields		3M Yields	
	Levels	Changes	Levels	Changes
1st PC % Variance Explained	0.91	0.66	0.88	0.70
2nd PC % Variance Explained	0.06	0.09	0.05	0.07
Average Pairwise Correlation	0.90	0.62	0.86	0.64
CD Test Statistic	201.8***	138.54***	194.16***	144.40***

Principal Component Analysis estimated after centering and standardizing variables to unit variance. Data consists of monthly long-term (10Y) and short-term (3M) bond yields across 20 countries from January 1999 to February 2021. The CD test statistic refers to the Pesaran [2021] test for cross-section dependence in panel data where the null hypothesis is no cross-sectional dependence.

proposed factor regression to test factor strength in long-term bond yields is:

$$y^{10Y}_{i,t} = a_i + B_{1i} y^{10Y}_{i,t-1} + B_{2i} y^{3M}_{i,t} + \pi_i \bar{y}^{10Y}_{-i,t} + e_{i,t}, \tag{A.1}$$

where country i 10-year (10Y or 'long-term') yields are regressed on its own lag and the country i 3-month (3M or 'short-term') yield as controls, along with a simple estimate of the global yield factor constructed as the equal-weighted cross-section average of 10-year yields from the other 19 advanced economies *excluding* country i (denoted $\bar{y}^{10Y}_{-i,t}$). The coefficient of interest is π_i, the factor loading on the global yield. Factor strength is determined by the proportion of π_i estimates which are statistically significant, out of the total 20 estimates (U.S. plus 19 additional countries).

Critical values which indicate significance are adjusted to account for multiple testing using the Bonferroni correction.[A1] At the 10% level, all but one estimate of π_i across the 20 country regressions are statistically significant (95%). At the 5% level, 18 out of 20 (90%) enter significantly, and at the 1% level, 16 of the 20 (80%) factor loading estimates are significant. Bailey et al. [2020] suggests that factors are considered weak if the proportion of significant factor loadings lies between 0 and 50%, semi-strong if they are above 50% but below 100%, and strong when 100% of factor loadings test as significant. Based on

[A1] The Bonferroni adjustment can be considered a standard case of the more general multiple testing correction described in Chudik et al. [2018].

these results, the global factor in long-term bond yields is confirmed as quite pervasive – and would likely enter even stronger in the absence of controlling for lagged 10-year yields and the short-term yield. As such, the factor in long-term bond yields may be considered a strong/semi-strong factor.

A2.1 Common Fo Yields and U.S. Yield Curve Factors

U.S. yield interest rate movements are explained well by small number of factors, usually three, known as level, slope, and curvature [Litterman and Scheinkman, 1991]. These factors are unobserved but typically extracted from the entire term structure of U.S. interest rates. An important question in the context of this study then is whether global factors help explain U.S. yields beyond the traditional yield curve factors. In other words, we ask whether the global yield factors are spanned by the domestic yield curve factors.

We estimate the level, slope, and curvature factors using the standard PCA approach on monthly U.S. interest rates across 10 maturities: 3-month, 6-month, 1-year, 2-year, 3-year, 5-year, 7-year, 10-year, 20-year, 30-year. The three factors generally explain over 99% of the variation in the U.S. yield curve, and my findings confirm that result. To test whether the global yield factors have any additional explanatory power over the yield curve factors for long and short-term U.S. interest rates, we consider the following separate regressions for the 10Y and 3M U.S. yield:

$$y_{us,t}^{10Y} = \beta_0^{10Y} + \beta_1^{10Y} level_{us,t} + \beta_2^{10Y} slope_{us,t} + \beta_3^{10Y} curve_{us,t} + \beta_4^{10Y} \mathcal{Y}_{g,t}^{10Y} + e_{us,t} \tag{A.2}$$

$$y_{us,t}^{3M} = \beta_0^{3M} + \beta_1^{3M} level_{us,t} + \beta_2^{3M} slope_{us,t} + \beta_3^{3M} curve_{us,t} + \beta_4^{3M} \mathcal{Y}_{g,t}^{3M} + e_{us,t}, \tag{A.3}$$

where the yield curve factors are given by $level_{us,t}$ for the level factor, $slope_{us,t}$ for the slope factor, and $curve_{us,t}$ for the curvature factor. The global yield factor for long and short-term interest rates are denoted $\mathcal{Y}_{g,t}^{10Y}$ and $\mathcal{Y}_{g,t}^{3M}$, respectively. A significant coefficient on β_4 would suggest that the global yield factors are unspanned by traditional U.S. yield curve factors, and that the international dimension is important to take into account when explaining yield variations in the United States. For each regression (A.2) and (A.3), the yield curve factors are estimated in slightly different ways to avoid collinearity issues during the regression estimation. For (A.2) the 10-year U.S. yield is excluded when extracting the factors from the yield curve, and similarly for (A.3) the 3-month U.S. yield is excluded when extracting the factors because these are the specific maturities we are trying to explain.[A2]

Table A.4: U.S. Yield Curve Factor Regressions

	Dependent Variable			
	10Y U.S. Yield		3M U.S. Yield	
Intercept	3.307***	(0.041)	1.453***	(0.106)
Level	1.284***	(0.018)	1.356***	(0.090)
Slope	0.414***	(0.015)	0.840***	(0.023)
Curvature	0.038***	(0.008)	0.125***	(0.033)
10Y Global Yield	0.035**	(0.015)		
3M Global Yield			0.381***	(0.073)
Adj. R^2	0.997		0.983	
T	266		266	

Regression estimates from Equations A.2 and A.3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Yield curve factors (level, slope, curvature) are standardized to zero mean and unit variance while the global yield factors are in yield levels. Level, slope, and curvature factors are extracted using Principal Component Analysis over the term structure of U.S. yields, and map to the first three principal components. For Equation A.2 the 10-year U.S. yield is excluded when extracting the factors from the yield curve, and similarly for Equation A.3 the 3-month U.S. yield is excluded when extracting the factors. Sample period: January 1999 to February 2021.

Table A.4 reports ordinary least squares (OLS) estimates for (A.2) and (A.3). For both the U.S. 10Y and 3M yield, corresponding global yield factors $\mathcal{Y}_{g,t}^{10Y}$ and $\mathcal{Y}_{g,t}^{3M}$ enter the regressions significantly after controlling for domestic U.S. yield curve factors, despite the yield curve factors alone being able to explain over 99% and 96% of the variation in 10Y and 3M U.S. yields, respectively. This suggests that indeed, global yield factors contain information relevant for U.S. interest rates that is not spanned by traditional domestic yield curve factors. However, this analysis cannot make any statements on the direction of causality, and assumes that in equlibrium these individual U.S. yields are a function of domestic yield curve

[A2] The estimated factors under omission of the 10-year yield are nearly perfectly correlated with the estimated factors under omission of the 3-month yield, and there is little to no practical difference between the two variants.

and global factors. It is possible however for U.S. yields to spillover to foreign yields given the sizable role of the United States economy and financial sector and at the same time, for global yields can influence U.S. yields. Additionally, both global factors, $\mathcal{Y}_{g,t}^{10Y}$ and $\mathcal{Y}_{g,t}^{3M}$, can be included in the regressions (A.2) and (A.3). When this is done, both long and short-term global factors enter as highly significant (at the 1 percent level) in both the 10-year U.S. yield regression and 3-month U.S. yield regression, further supporting the role of global factors in explaining U.S. yields.

A3 Alternative Common Factor Weights

This section compares different approaches to recover bond yield factors from the cross-section of advanced economy government bond yields. The benchmark approach applies GDP weights, which vary over time and take into account economic size differences across countries. For robustness, alternative approaches considered include public debt weights, which also vary over time but assign weights based on the U.S. Dollar value of the stock of national public debt. Debt-based weights, for instance, make Japan's bond yields much more influential in the factor estimation given the size of their public debt market even relative to their GDP. Statistical approaches are also considered. First, a cross-section average (CSA) approach which is the simple cross-section equal-weighted average each period across the 19 non-U.S. country yields. Finally, we also estimate a PCA-based factor which assigns country weights based on estimated PCA factor loadings. Note that the statistical approaches assign time-invariant weights across bond yields.

Tables A.5 and A.6 report correlation coefficients between the different variants of the global long-term short-term and yield factors, respectively. The correlations are very high, most above 0.99 which implies that the global yield factors are relatively robust to alternative weight schemes. The reason for this is because of the high pair-wise correlations across international bond yields to begin with – as individual interest rates become increasingly

Table A.5: Global 10-Year Yield Factor Correlations Across Alternative Weight Schemes

	Debt	CSA	PCA
GDP	0.996	0.999	0.999
Debt		0.993	0.995
CSA			1.000

Correlations between global long term bond yield factors constructed under varying weighted averages of 19 non-U.S. 10-year government yields (Equation 2). GDP refers to the benchmark GDP-weighted average. Debt refers to national public debt weighted average of international bond yields. CSA refers to the simple equal-weighted cross-section average. PCA refers to weights assigned via Principal Component Analysis. Estimation period covers January 1999 to February 2021.

Table A.6: Global 3-Month Yield Factor Correlations Across Alternative Weight Schemes

	Debt	CSA	PCA
GDP	0.994	0.999	0.999
Debt		0.991	0.993
CSA			1.000

Correlations between global short term bond yield factors constructed under varying weighted averages of 19 non-U.S. 10-year government yields (Equation 2). GDP refers to the benchmark GDP-weighted average. Debt refers to national public debt weighted average of international bond yields. CSA refers to the simple equal-weighted cross-section average. PCA refers to weights assigned via Principal Component Analysis. Estimation period covers January 1999 to February 2021.

synchronized across countries, alternative weighted-average composites of the yields also become increasingly synchronized.

A4 Common Foreign Yield Factors, Global Growth, and UST Flows

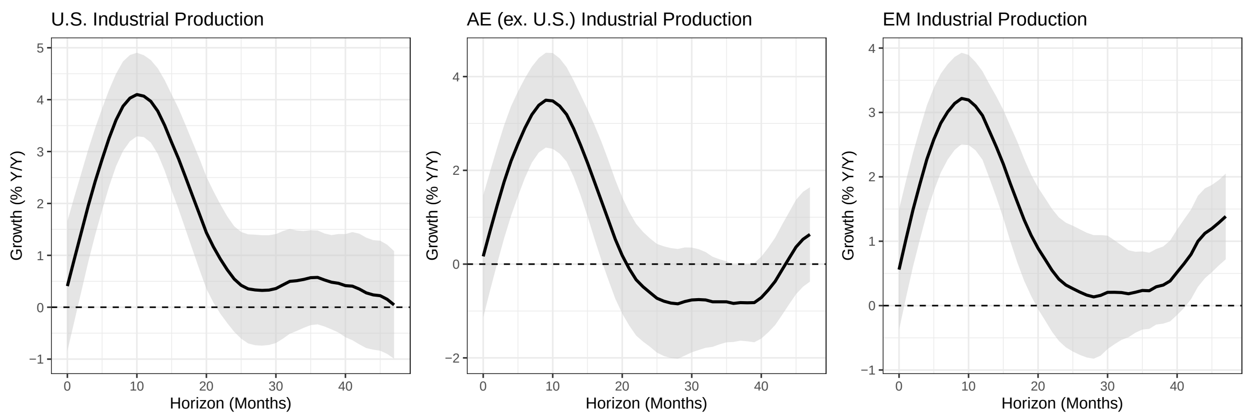
In Figures A.1 and A.2, we provide additional evidence consistent with previous studies finding that global yields are indeed positively associated with broad economic expansions. Implementing simple local projection regressions of the form:

$$Y_{t+h} = \alpha^h + \beta_1^h \mathcal{Y}_{g,t}^{10Y} + \beta_2^h \mathcal{Y}_{g,t}^{3M} + e_t^h, \tag{A.4}$$

Figure A.1 plots the estimated coefficient of β_1^h which corresponds to the association of different outcomes (Y_{t+h}) with a 1 percentage point increase in the global long-term yield factor, $\mathcal{Y}^{10Y}_{g,t}$ at different horizons. For monthly frequency data, horizon h spans 48 months, from contemporaneous ($h = 0$) to $h = 47$ months ahead. For the quarterly frequency data, horizons range from 0 to 23 quarters for a total of 24 quarters.

Figure A.1 shows that an increase in long-term global yields ($\mathcal{Y}^{10Y}_{g,t}$) is associated with economic expansions across the U.S., advanced, and emerging economies.[A3] Note that economic activity responds with a lag to changing global yields, highlighting the forward-looking nature of the latter and the importance of using yields to construct the global factor governing U.S. Treasuries.

Figure A.1: Long-Term Global Yields and their Correlation with Global Growth



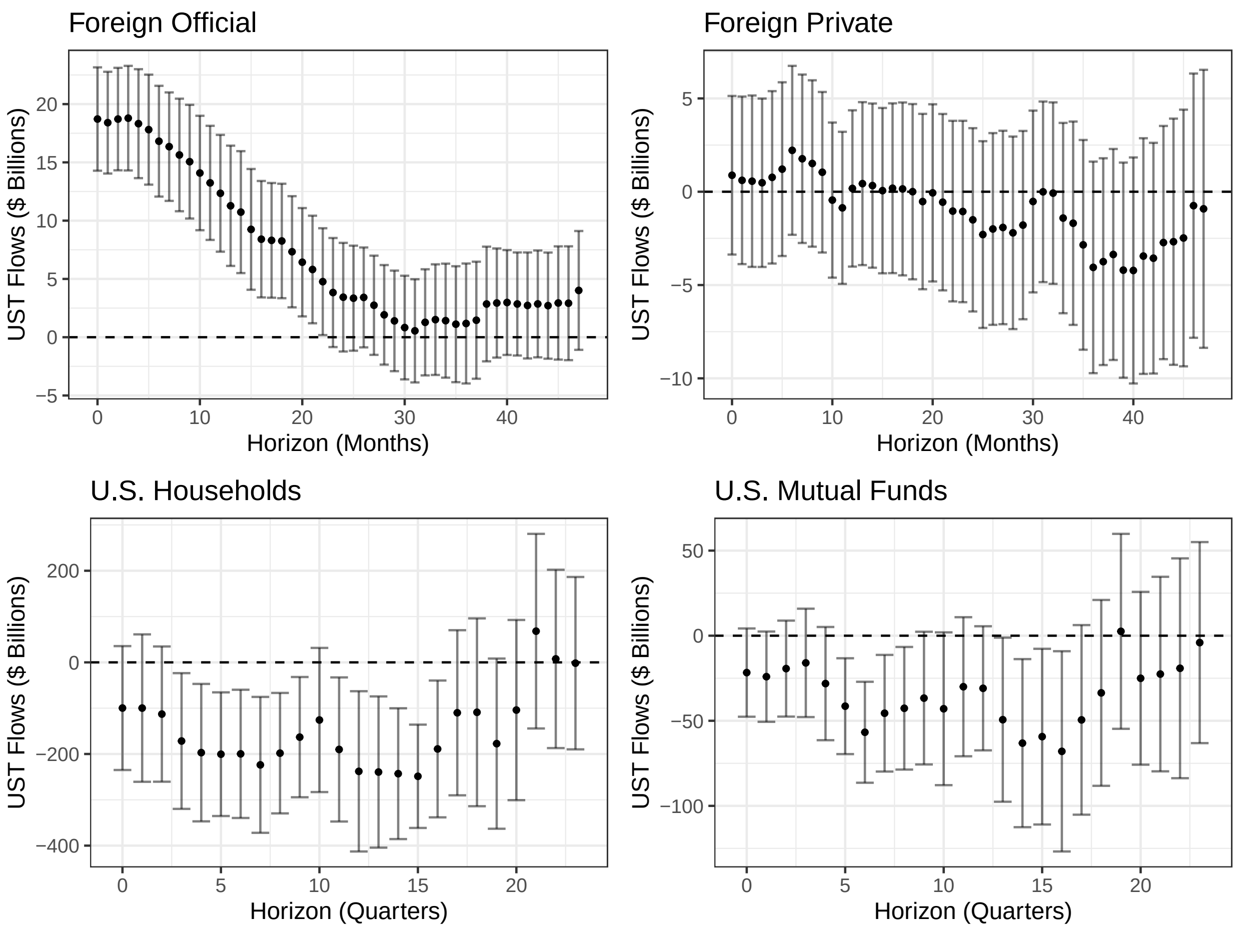
Global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2. Dynamic associations estimated from Equation A.4. Sample period is January 1999 to December 2018. All variables are monthly frequency.

Figure A.2 shows that when global yields rise (and as the world economy expands), foreign officials purchase more USTs. By contrast, a similar pattern does not appear among foreign private UST flows. Moreover, U.S. households and mutual funds sell USTs when global yields are rising and buy them in a flight to safety when global yields fall amid slowing global growth.[A4] Taken together, we see that global yields reflect global economic conditions, and

[A3] These are monthly industrial production growth measures taken from the Dallas Fed Database of Global Economic Indicators, which covers a core sample of 40 countries.

[A4] Household (which includes hedge funds) and mutual fund UST transactions data are quarterly frequency and taken from the Federal Reserve Financial Accounts.

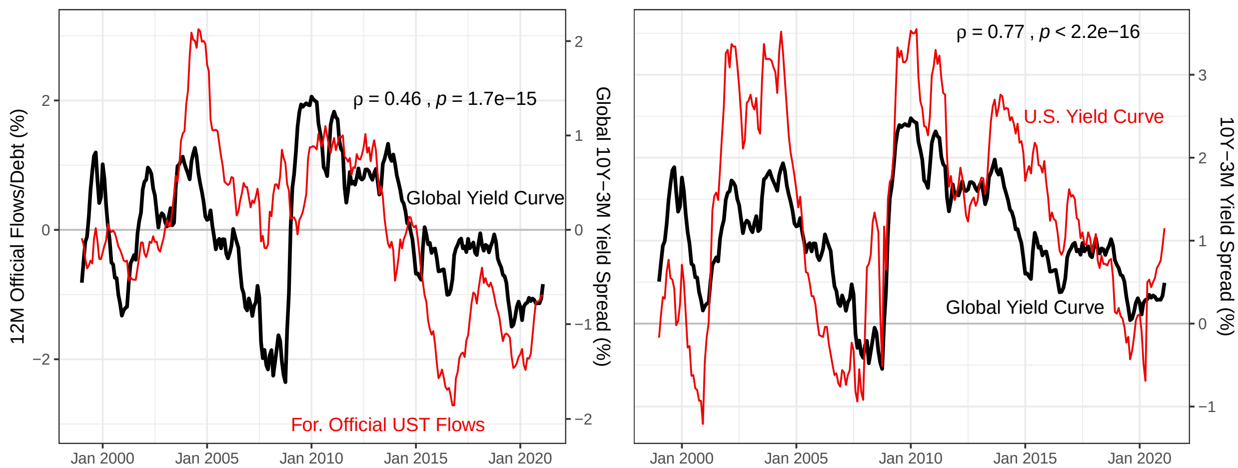
Figure A.2: Long-Term Global Yields and their Correlation with Foreign and Domestic UST Demand



Global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2. Dynamic associations estimated from Equation A.4. Foreign official and private flow data uses adjusted benchmark-consistent flows Bertaut and Tryon [2007] and Bertaut and Judson [2014]. Sample period is January 1999 to December 2018. Foreign UST flows are monthly frequency and U.S. Households and U.S. Mutual Fund UST flows which are quarterly.

these conditions shape the appetite for USTs across different investors. Uniquely, foreign official demand for USTs is strongly pro-cyclical, while the UST demand from traditional investors is counter-cyclical or acyclical.

Figure A.3: The Global non-U.S. Yield Curve with Foreign Official Net UST Purchases (left) and the U.S. Yield Curve (right)



The thick line is the global 10-year less 3-month yield spread. Global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2. Left panel overlays the global yield curve with TIC-reported foreign official UST flows data, and both series are de-meaned. Right panel overlays the global yield curve with the 10-year less 3-month U.S. yield curve, and series are not de-meaned. Foreign official UST flows are calculated as a 12-month rolling sum.

A4.1 Foreign Official UST Flows and Common Foreign Yield Factor

The joint effect of a global factor on U.S. yields and foreign demand for Treasuries will bias the observed relationship between U.S. yields and official UST demand upward if not accounted for. For instance, the simple correlation between U.S. 10Y yields (10-year minus 3-month yield spread) and official UST flows-to-marketable debt[A5] over the sample period is 0.22 (0.32). The correlation between the global 10-year yield factor (global 10-year minus global 3-month yield spread) and official UST flows is 0.42 (0.36). Both are positive, reflecting the

[A5]Marketable debt is computed as total U.S. federal debt outstanding less debt held by Federal Reserve banks.

cyclical nature of foreign official UST flows which masks the intuitively negative effect that exogenous UST demand shocks should have on U.S. yields. Rather, the positive correlations are consistent with global expansions inducing international reserve accumulation and higher interest rates jointly. Figure 3 plots net purchases of USTs by foreign officials, scaled by marketable U.S. debt outstanding using both raw TIC and benchmark-consistent flow data of Bertaut and Tryon [2007] and Bertaut and Judson [2014].[A6] Figure A.3 traces the global (non-U.S.) 10Y-3M yield curve measured as $\mathcal{Y}_{g,t}^{10Y} - \mathcal{Y}_{g,t}^{3M}$ as a proxy for the global economic cycle. The left-panel shows a clear positive relationship between the global yield curve and 12-month net foreign official UST purchases. The right-panel shows that the U.S. yield curve ($y_{us,t}^{10Y} - y_{us,t}^{3M}$) covaries strongly with the global yield curve as well.

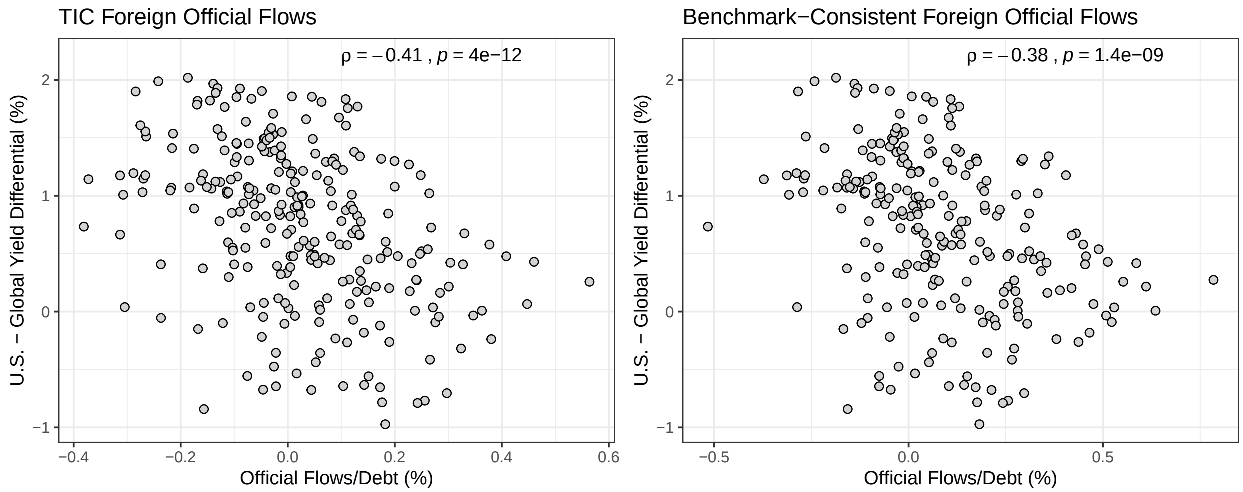
A4.2 Foreign official UST flows are negatively associated with the U.S.-global yield differential

While the correlation between global yields and official UST flows is positive (pro-cyclical UST demand by foreign officials), the correlation between official UST flows and the 10Y U.S.-global yield *differential* is significantly negative (-0.41) as shown in Figure A.4 when using either raw TIC flows data or the benchmark-consistent flows data. The correlation becomes increasingly negative, reaching -0.52 when using 12-month rolling official flows instead of monthly flows (Figure A.5). The yield differential is a simple yet perhaps crude way to partial out the component of U.S. yields driven by global factors like world economic conditions and neutral rates. This reveals a theoretically consistent negative relationship between foreign official UST purchases and U.S. yields.

Any impact should be in excess of the movement in global non-U.S. yields so long as UST purchases disproportionately impact U.S. yields relative to yields of other countries. The significant negative correlation suggests that foreign official UST purchases put downward pressure on U.S. 10Y yields relative to non-U.S. 10Y yields. The negative correlation is

[A6] Detail on benchmark-consistent flows are provided in the following section.

Figure A.4: Foreign Official Net UST Purchases and U.S.-Global 10Y Yield Differentials

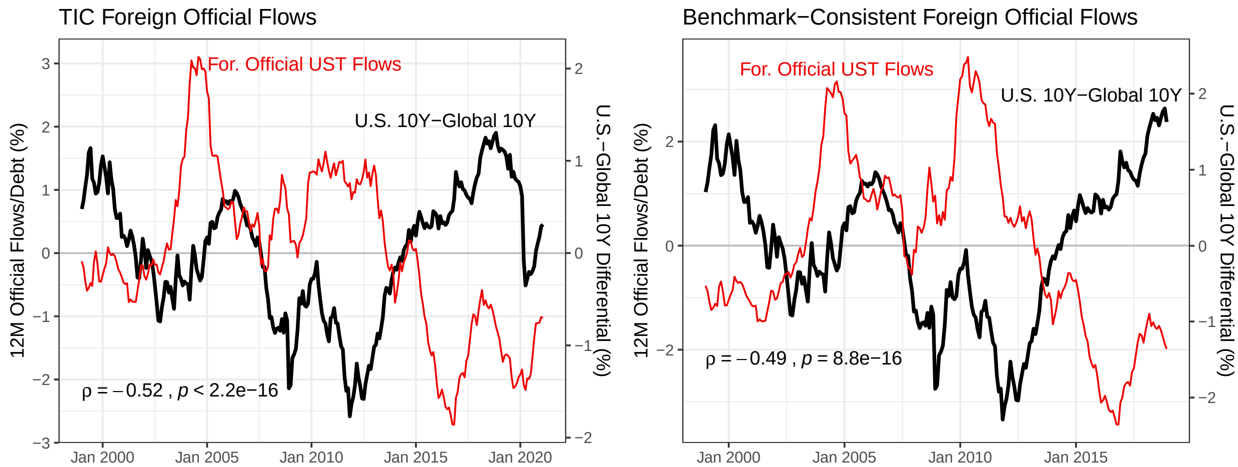


X-axes are monthly purchases/sales of UST by foreign officials scaled by 12-month lagged U.S. marketable debt. Y-axes are 10-year U.S. - global yield differentials, where global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2. Left panel uses TIC-reported flows data (January 1999 to February 2021) on the x-axis and right panel uses adjusted benchmark-consistent flows on the x-axis following Bertaut and Tryon [2007] and Bertaut and Judson [2014] (January 1999 to December 2018).

also consistent with inelastic foreign official demand for USTs, i.e. causality running from UST purchases to U.S. yields, at least contemporaneously. If U.S. yield differentials were causing foreign official flows to respond, one would expect the correlation in Figure A.4 to be positive under uncovered interest rate parity (UIP) as wider and more positive U.S. interest rate spreads attract more UST inflows from abroad. However, an alternative explanation for this negative relationship may toward causality from U.S. yields to official UST flows, is such that US monetary policy divergence from the rest of the world induces foreign currency depreciation, prompting reserve managers to intervene and sell their UST holdings.

The relationship between foreign official flows and the U.S.-global yield differential may also be dynamic, as any evidence of foreign official flows forecasting U.S. yields (or yield differentials) would be informative about the direction of causality between flows and yields. As a simple empirical test to capture dynamic correlations, we extend the static correlation analysis to a dynamic setting by estimating local projections as in Jordà [2005] in Section A4.3 of the Appendix. Foreign official flows are significantly correlated with future changes

Figure A.5: Time Series of Foreign Official Net UST Purchases and U.S.-Global 10Y Yield Differentials



The thick line is the U.S. 10-year yield - global 10-year yield differential. Global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2. The thin line refers to foreign official UST flows data, and both series are de-meaned. Foreign official UST flows are calculated as a 12-month rolling sum. Left panel uses TIC-reported flows data (January 1999 to February 2021) on the x-axis and right panel uses adjusted benchmark-consistent flows on the x-axis following Bertaut and Tryon [2007] and Bertaut and Judson [2014] (January 1999 to December 2018).

in the U.S.-global 10-year yield differential.

A4.3 Local projections of the U.S.-global yield differential

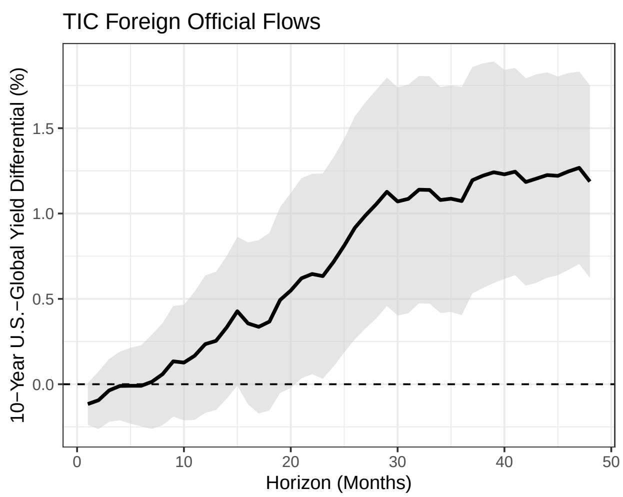
The negative relationship between foreign official flows and the U.S.-global yield differential shown in Figure A.4 may also be dynamic, as any evidence of foreign official flows forecasting U.S. yields (or yield differentials) would be informative about the direction of causality between flows and yields. As a simple empirical test to explore dynamic correlations, we extend the static correlation analysis to a dynamic setting by estimating local projects as in Jordà [2005] of the following form:

$$y^{10Y}_{us,t+h} - \mathcal{Y}^{10Y}_{g,t+h} = \alpha^h + \beta^h_1[y^{10Y}_{us,t} - \mathcal{Y}^{10Y}_{g,t}] + \beta^h_2 \Delta FO_t + e_t, \tag{A.5}$$

regressing the h-month ahead 10-year U.S.-global yield differential denoted $y^{10Y}_{us,t+h} - \mathcal{Y}^{10Y}_{g,t+h}$, on the month t yield differential and foreign official purchases and sales of U.S.

Treasury bonds and notes (scaled by 12-month lagged U.S. marketable debt), denoted ΔFO_t. Local projection impulse responses are composed of the sequence of $\hat{\beta}_2^h$ estimates for $h = 1, ..., 48$ months. These local projection estimates are recovered using the standard procedure of estimating $H = 48$ separate regressions. The impulse response traced in Figure A.6 then shows the forward evolution of the 10-year U.S.-global yield differential following a 1 percentage point decrease or sale of foreign official USTs (as a percentage of GDP).

Figure A.6: Local Projection Impulse Response of U.S.-Global 10Y Interest Rate Differentials to a Foreign Official UST Sale (1% of Debt)



Y-axes are the response of 10-year U.S. - global yield differential in percent to a 1-percentage point drop in foreign official UST flows scaled by 12-month lagged U.S. debt outstanding less debt held by Federal Reserve banks, where global yields are constructed as GDP-weighted averages of non-U.S. yields as in Equation 2. Estimates are recovered from Equation A.5 via local projections. Shaded region refers to 90% confidence interval based on standard errors that are adjusted for heteroscedasticity and autocorrelation. Left panel uses TIC-reported flows data (January 1999 to February 2021) and right panel uses adjusted benchmark-consistent flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] (January 1999 to December 2018).

The left panel uses TIC raw flows data while the right panel uses benchmark-consistent flows. The local projections show that UST sales (purchases) by foreign officials put significant upward (downward) pressure on U.S. long-term yields relative to global long-term yields. Following a foreign official UST sale amounting to 1 percentage point of marketable debt, U.S. 10-year yields widen relative to rest-of-world 10-year yields by roughly 100 basis points at the peak (roughly 55 basis points per $100 billion flow based on marketable debt of $18 trillion). Since currently no other factors are being controlled for, these statistics should be interpreted as dynamic correlations only.

A5 Robustness of Regression-Based Estimates

A5.1 Details on the regression specification and estimation

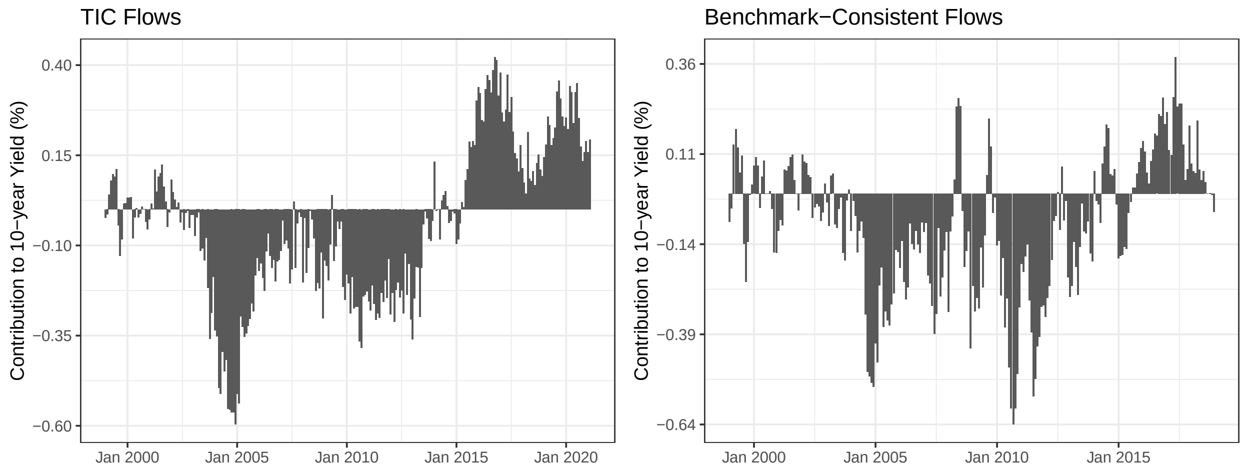
There are a few notable features of the model specification (3) that are worth discussing. First, the inclusion of a lagged dependent variable has been a regular subject of debate.[A7] While including the lag helps ensure stationary of the time-series and purges the residuals of serial correlation, it biases the OLS estimates in short samples and may absorb important variation thats not necessarily attributed to past values of the dependent variable under the true model. Rather, the inclusion of a lagged dependent variable may lead to a misspecified model. Alternatively, if the 10-year yield and other covariates are cointegrated, estimating (3) in levels gives (super)consistent estimates and the lagged term is not necessary – and in fact should not be included in the regressor if it is not a part of the true model. Theory guides us to believe that interest rates are indeed cointegrated as the U.S. long-term yield cannot deviate too far from the short-term yield, or from comparable long-term yields of other countries without arbitragers intervening. Theoretical guidance along with previous empirical evidence helps, to an extent, justify a specification in levels without a lagged dependent variable. However, absent the lagged dependent variable the regression residuals are still likely to exhibit significant serial correlation and appropriate standard errors which correct for this must be applied. Because we cannot observe the true underlying model and there is no consensus on whether or not to include a lagged dependent variable in this setting, we will estimate versions of (3) without and with a lagged dependent variable, the former case will act as the baseline and imposes the restriction $\phi_0 = 0$. In all sets of regression results without the lagged dependent variable, statistical tests strongly suggest that the

[A7] Warnock and Warnock [2009] does not include one, Bertaut et al. [2012] does, while Beltran et al. [2013] estimates a regression in first-differences, which essentially includes a lagged dependent variable but imposes that the coefficient $\phi_0 = 1$.

10-year yield is cointegrated and residuals are stationary.

The specification differs from the literature by including L lags of foreign official UST purchases. The literature usually considers a 12-month rolling sum of foreign UST purchases as the main covariate of interest, with the coefficient interpreted as a 'long-run flow effect'. However, this is a special case of (3) where $L = 11$ and $\theta_l = \theta$ for all $l = 0, \ldots 11$. Therefore (3) generalizes the previous modeling approaches by allowing for the possibility of different coefficients on each lag of ΔFO_t. The estimated cumulative or long-run impact of UST purchases or sales on U.S. yields over 12 months is given by $\sum_{l=1}^{11} \theta_l/(1-\phi_0)$. In the case where $\phi_0 = 0$ (excluding the lagged dependent variable), this is just the sum of coefficients across all L lags of ΔFO_t plus the contemporaneous flows, and computing standard errors for the sum of coefficients is straightforward. When we include the lagged dependent variable ($\phi_0 \neq 0$), we must take into account the feedback effects of ΔFO_{t-l} on y_t^{10Y} through y_{t-1}^{10Y}. This overall effect is consistently estimated by dividing the sum of coefficients θ_l by $1 - \phi_0$ as shown in standard autoregressive distributed lag (ARDL) models and done in Bertaut et al. [2012]. Because the total effect of official flows in the presence of a lagged dependent variable is now a non-linear combination of regression coefficients, the standard errors for the 12-month overall impact of flows on Treasury yields are computed via bootstrap procedures. Specifically, we estimate the regression model, and then bootstrap the model residuals, sampling with replacement 1,000 times to recover the simulated distribution of $\sum_{l=1}^{11} \theta_l/(1-\phi_0)$ from which we compute standard errors.[A8] All reported standard errors are adjusted for heteroscedasticity and autocorrelation using the standard Newey-West adjustment [Newey and West, 1987] with one lag.

Figure A.7: Estimated Historical Contribution of 12-month Foreign Official UST Flows to 10-Year U.S. Yields



Contributions derived from estimates of Equation 3. Left panel is the TIC-reported flows data and right panel are adjusted benchmark-consistent flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014].

A5.2 Historical Decomposition of Foreign Official UST Flow Impact on 10-year U.S. Yields

Figure A.7 traces the historical contribution of foreign official UST flows (as a percentage of marketable debt) on the U.S. 10-year yield over the sample period (through February 2021 for TIC flows and through December 2018 for benchmark-consistent flows). This decomposition is based on the estimates from baseline regression (3). The results confirm previous studies which document substantial yield compression attributed to foreign demand for USTs. In 2005, U.S. 10-year yields were lower by roughly 50 basis points compared to the scenario of zero foreign official flows. Again in 2011 foreign official flows contributed to substantial compression in U.S. yields: up to -60 basis points based on benchmark-consistent flows. The impact of foreign flows on yields peaked around September 2011 during the U.S. debt ceiling crisis. Ironically, the U.S. debt crisis sparked an international flight-to-safety, as investors

[A8] Alternatively for non-linear combinations of regression coefficients, standard errors can be computed using the delta method. However, the delta method specifically calculates standard errors for a linear approximation of the non-linear function of coefficients, and this approximation may or may not be reliable.

sought the safety of U.S. Treasuries. Since 2015, however, there's been an increasing shift in persistent foreign official UST net sales rather than purchases. As a result, foreign official UST flows have been putting upward pressure on U.S. yields since 2015 which coincides with the year China's Renminbi entered the IMF's reserve currency basket and the country renewed aims to foster greater exchange rate flexibility. Since then, foreign official UST sales may have pushed U.S. yields higher by roughly 40 basis points compared to a the 'zero foreign official flow' case.

A5.3 Estimates from an Error-Correction Model

Another popular econometric model used to estimate the impact of foreign official UST demand on US yields is the error correction model (ECM) which disentangles both short-run and long-run effects in the presence of cointegration [Bandholz et al., 2009; Warnock and Warnock, 2009; Beltran et al., 2013]. To test the robustness of the regression specification, this section estimates a simple first-order ECM for U.S. 10-year yield changes:

$$\Delta y^{10Y}_{us,t} = \mathbf{\Phi} \begin{bmatrix} \Delta y^{10Y}_{us,t-1} \\ \Delta y^{3M}_{us,t-1} \\ \Delta \mathcal{Y}^{3M}_{g,t-1} \\ \Delta \mathcal{Y}^{10Y}_{g,t-1} \\ \Delta FO_{t-1} \\ \Delta \mathbf{X}_{t-1} \end{bmatrix} + \Gamma[\hat{\epsilon}_{us,t-1}] + u_{us,t}, \tag{A.6}$$

where the second term on the right-hand side, $\hat{\epsilon}_{us,t-1}$, corresponds to the lagged residual from (3) which estimated the cointegrating relationship for 10-year U.S. yields in levels. We can recover the error-correcting estimate of official flows on short-run (1-month ahead) changes in U.S. yields by taking the product of -1, the long-run effects previously estimated as $\sum_{l=0}^{11} \theta_l/(1-\phi_0)$, and the estimate of Γ. Specification (A.6) is estimated via OLS with an intercept term but no linear time trend, and standard errors are adjusted for autocorrelation and heteroscedasticity.

Under the baseline specification of the cointegrating relationship (3) with no lagged dependent variable and under benchmark-consistent foreign official flows scaled by U.S. marketable debt, Γ is estimated to equal -0.137 with a standard error of 0.045, significant at the 1% level. This implies that a foreign official purchase amounting to 1 percentage point of marketable debt reduces 10-year yield the following month by 15 basis points ($-1 \times -1.105 \times -0.137 = -0.154$) while a \$100 billion sale would push U.S yields down by about 8.5 basis points. These estimates are in line with the regression estimates using the stock of foreign official UST holdings instead of 12-month flows, as well the as error-correction based estimates of Beltran et al. [2013].

If global yield factors were not considered in this procedure, the estimate of Γ is equal to -0.159, implying that a foreign official purchase amounting to 1 percentage point of marketable debt reduces 10-year yield the following month by 5.5 basis points ($-1 \times -0.348 \times -0.159 = -0.0588$) while a \$100 billion sale would push U.S yields down by about 3 basis points. Effect sizes are roughly two to three times larger after accounting for global factors.

A5.4 Estimates under alternative regression specifications

For robustness, the large changes in estimated effect sizes upon introducing global factors as controls are similarly found in specifications when assigning the term premium as the dependent variable (Table A.7), that include a lagged dependent variable (Table A.8), and when scaling foreign official flows by U.S. GDP instead of marketable debt (Table A.9) as done in Warnock and Warnock [2009]. In both the case of including a lagged dependent variable and scaling by GDP, the absolute size of the effect is larger than those reported in the baseline results which do not include a lagged dependent variable and scale flows by marketable debt. Table A.10 extends the sample to start from 1990 and finds that the results are generally robust to the lengthening the sample period. The estimated effect sizes do, however, decrease as the 1990s were characterized by relatively benign foreign official

UST holdings and flows, possibly indicating the presence of a structural break in the time-series relationship after foreign officials became an increasingly dominant participant in US Treasury markets. We also consider regression specifications that replace 12-month UST flow with the level of foreign official UST holdings, to estimate the permanent or long-run 'stock effect' (Table A.11). Consistent with Beltran et al. [2013], the permanent effect of a change in foreign official UST holdings is quantitatively smaller than the 12-month flow effect.

A5.5 Including additional global factors

Table A.12 re-estimates the baseline regression (3) with long and short-term yield factors, and reports the long-run estimated impact of foreign official flows on U.S. 10-year yields. Different weight schemes are applied to construct the global factors, with the GDP-weighted factor serving as the baseline (same as Table 1). Regardless of weighting approach, estimates of the impact of foreign official flows on U.S. yields remains highly significant and negative after introducing global factors. It may be worth noting that under debt-based weights which put heavy emphasis on Japanese yields, the size of the effect is estimated to be even larger in absolute terms, while the estimated impact slightly falls under CSA and PCA based weights. Based on these alternative weight schemes along with the baseline GDP weight, a $100 billion sale of USTs by foreign officials, assuming $18 trillion marketable debt, would push 10-year yields up between 45 and 76.23 basis points over the long-run (referring to benchmark-consistent flows).

The following three rows of Table A.12 refer to estimates when using global factors which try to correct for the two-way feedback effects between U.S. yields and foreign yields. The first two of three approaches considered are 1) using lagged global factors, 2) using the residual from a regression of the 10-year global yield factor on the U.S. 10-year yield and 3-month global yield factor on the U.S. 3-month yield in order to purge any contemporaneous correlation between U.S. 10-year yields and foreign 10-year yields. Referring to the benchmark-consistent foreign official flow measure, a 1 percentage point drop in flows relative to marketable debt is associated with a 103.7 basis point rise under the lagged specification, and a 47.8 basis point rise under the residual specification. Finally, we also consider a modified regression where the 10-year U.S. yield is replaced with the 10-year U.S.-global yield spread as the dependent variable (and the global 10-year yield is removed from the set of conditioning variables) in order to focus attention on simply explaining the global 10-year spread rather than the 10-year U.S. yield. The results show that the same 1 percentage point change in flows is associated with a 144.8 basis point change in the 10-year U.S.-global spread.

The final set of rows in Table A.12 augments the baseline regression with additional conditioning variables to capture the presence of additional global factors. First we consider the G10 Treasury premium, or average CIP deviations for both 3-month and 10-year cross-country yields. Covered interest rate parity (CIP) implies that the yield offered on U.S. risk-free Treasury bonds of a given maturity should be equivalent to those of risk-free bonds of a foreign government of the same maturity whose foreign currency cashflows are swapped into U.S. Dollars. However, Du and Schreger [2016] and Du et al. [2018] document persistent differences between these two U.S. Dollar based yields of risk-free government bonds. This deviation is denoted as the U.S. Treasury premium, measuring the premium attached to the special liquidity and safety features of U.S. Treasury bonds over their synthetic counterparts. Using Treasury premiums for the G10 countries provided by Du and Schreger [2016] and Du et al. [2018], U.S. average Treasury premia are computed for both 3-month and 10-year tenors. The individual G10 U.S. Dollar Treasury premia averaged over are: Australia, Canada, Switzerland, Denmark, Euro, United Kingdom, Japan, Norway, Denmark, Euro Area, Norway, New Zealand, and Sweden.[A9]. The estimated long-run impact of foreign official UST flows on U.S. 10-year yields falls slightly but remains large and significant. Referring to the benchmark-consistent flow measures, a fall in flows relative to marketable

[A9] For months where data are unavailabe for any given country, the average is computed over all other countries with available data.

debt of 1 percent is associated with a rise in U.S. yields of 82.7 basis points (45.4 basis points per $100 billion flow) after controlling for U.S. Treasury premia.

Geopolitical risk may also jointly affect the incentive to hoard international reserves by foreign officials, while also affecting U.S. yields. Considering the news-based measure of geopolitical risk by Caldara and Iacoviello [2018], we find that the long-run flow effect of foreign official UST demand on U.S. yields decreases slightly, but remains large and significant after accounting for geopolitical risk.

While the forward-looking nature of U.S. long-term yields is difficult to capture, the baseline specification aims to account for this feature by including measures of future expected GDP growth, inflation, and risk premia (VIX). However, this may be an imperfect solution. To sharpen the analysis, we also consider a new measure of Fed forward guidance shocks of Swanson [2021] recovered from high-frequency asset price co-movements around FOMC announcements. These shocks are aggregated to the monthly frequency as daily sums, and two measures enter the baseline regression as additional controls: the monthly shocks themselves, along with the cumulative sum of monthly shocks. The latter measure is important to account for the fact that the dependent variable is the U.S. yield level, rather than yield change. Table A.12 shows that the long-run flow impact of foreign official flows on U.S. yields remains nearly unchanged after including additional forward guidance controls.

Fed forward guidance, by affecting the expected path of monetary policy is likely to affect 10-year U.S. yields. Unconventional large scale asset purchase (LSAP) policies which directly target long-term Treasuries are also a plausible factor that needs to be accounted for, specifically because these policies tend to be cyclical and administered during global crises. Global crises periods tend to also be the episodes that see the largest foreign official UST liquidations. After controlling for both Fed forward guidance and LSAP, the estimated impact of foreign official UST demand on U.S. yields increases substantially. This result is consistent with the overall premise that the effects of foreign official UST flows on U.S. yields tend to be dampened because foreign officials buy when others tend to sell, and *vice versa*.

Similarly, these results suggest that when the Federal Reserve intervenes and buys bonds, the impact of foreign official UST sales are dampened, as foreign official selling pressure is met with the Fed's purchasing liquidity.

It may be useful to also include a measure of global real economic activity rather than relying on global yields which may not always be a good approximation for world economic conditions. Using a measure of global (ex. U.S.) industrial production (IP) growth taken from the Dallas Fed Database of Global Economic Indicators, including global IP growth as an additional global factor does not have any substantial affect on the results.

It may be the case that shocks specific to emerging markets should also be accounted for among global factors, in the rare case that emerging market shocks spillover to the United States. Emerging market financial conditions are controlled for using monthly changes in the EMBI sovereign spread index and results for long-run estimates are shown in Table A.12, which are close to the baseline results.

Table A.7: 10-Year U.S. Term Premium Regressions

	Dependent Variable: 10Y U.S. Term Premium							
	TIC Flows				Benchmark-Consistent Flows			
3M U.S. Yield	-0.097***	(0.020)	-0.081***	(0.017)	-0.029	(0.024)	-0.033	(0.021)
1Y GDP Forecast	-0.062	(0.043)	-0.034	(0.045)	0.174**	(0.081)	0.188**	(0.087)
10Y Inflation Forecast	0.171	(0.474)	-0.360	(0.343)	0.048	(0.466)	-0.254	(0.343)
1Y Inflation Forecast	0.019	(0.058)	0.034	(0.054)	-0.016	(0.050)	-0.046	(0.051)
VIX	-0.004	(0.005)	-0.003	(0.005)	0.014**	(0.006)	0.010	(0.006)
Budget Surplus	-0.041	(0.028)	-0.008	(0.024)	-0.029	(0.032)	-0.014	(0.031)
12M Foreign Official Flows	-0.719***	(0.088)	-1.407***	(0.059)	-0.249**	(0.113)	-0.764***	(0.095)
10Y Global Yield			0.474***	(0.050)			0.376***	(0.057)
3M Global Yield			-0.141***	(0.045)			-0.057	(0.048)
Adj. R^2	0.769		0.823		0.726		0.773	
T	266		266		240		240	
ADF Statistic	-4.812***		-4.250***		-5.125***		-4.607***	

Regression estimates from Equation 3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Coefficients and standard errors for foreign official flows are estimated over the sum of coefficients on ΔFO_{t-l} where $l = 0, ...11$ over the last 12 months. Foreign official flows variable is scaled by U.S. marketable debt lagged 12 months. Regressions include an intercept term and linear time trend. Augmented Dickey-Fuller (ADF) test on regression residuals reject the null hypothesis that residuals are nonstationary. The sample periods under TIC flows and benchmark-consistent flows are January 1999 to February 2021 and January 1999 to December 2018, respectively. Benchmark-consistent flows data are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

Table A.8: 10-Year U.S. Yield Regressions with Lagged Dependent Variable

	Dependent Variable: 10Y U.S. Yield							
	TIC Flows				Benchmark-Consistent Flows			
3M U.S. Yield	0.040***	(0.012)	0.090***	(0.017)	0.057***	(0.019)	0.105***	(0.023)
1Y GDP Forecast	-0.018	(0.025)	-0.010	(0.027)	-0.003	(0.055)	0.016	(0.053)
10Y Inflation Forecast	-0.180	(0.256)	-0.342	(0.235)	-0.496**	(0.248)	-0.532**	(0.244)
1Y Inflation Forecast	-0.029	(0.033)	0.008	(0.032)	-0.008	(0.030)	0.008	(0.030)
VIX	-0.009***	(0.003)	-0.008***	(0.003)	-0.008**	(0.003)	-0.006*	(0.004)
Budget Surplus	-0.012	(0.015)	0.002	(0.014)	-0.032**	(0.019)	-0.014	(0.019)
12M Foreign Official Flows	-0.445	(1.577)	-2.012**	(0.842)	-0.907	(1.189)	-1.550**	(0.671)
10Y Global Yield			0.260***	(0.050)			0.249***	(0.053)
3M Global Yield			-0.145***	(0.034)			-0.136***	(0.033)
Adj. R^2	0.980		0.983		0.977		0.980	
T	265		265		239		239	
ADF Statistic	-11.023***		-9.750		-11.033***		-9.755***	

Regression estimates from Equation 3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Coefficients and bootstrapped standard errors for foreign official flows are estimated over the sum of coefficients on ΔFO_{t-l} where $l = 0, ...11$ over the last 12 months divided by 1 minus the autoregressive coefficient. Bootstrapped standard errors for foreign official flows based on 1,000 bootstrap samples of model residuals. Foreign official flows variable is scaled by U.S. marketable debt lagged 12 months. Regressions include an intercept term, linear time trend, and lagged dependent variable. Augmented Dickey-Fuller (ADF) test on regression residuals reject the null hypothesis that residuals are nonstationary. The sample periods under TIC flows and benchmark-consistent flows are February 1999 to February 2021 and February 1999 to December 2018, respectively. Benchmark-consistent flows data are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

Table A.9: 10-Year U.S. Yield Regressions with Official UST Flows Scaled by GDP

	Dependent Variable: 10Y U.S. Yield							
	TIC Flows				Benchmark-Consistent Flows			
3M U.S. Yield	0.254***	(0.029)	0.303***	(0.022)	0.363***	(0.033)	0.379***	(0.026)
1Y GDP Forecast	0.011	(0.066)	0.038	(0.064)	0.480***	(0.102)	0.445***	(0.102)
10Y Inflation Forecast	0.046	(0.661)	-0.725*	(0.397)	0.253	(0.610)	-0.324	(0.390)
1Y Inflation Forecast	0.057	(0.077)	0.142**	(0.065)	-0.044	(0.066)	-0.035	(0.056)
VIX	-0.021***	(0.004)	-0.016***	(0.005)	0.010	(0.006)	0.004	(0.006)
Budget Surplus	-0.081***	(0.036)	-0.022	(0.025)	-0.052	(0.041)	-0.038	(0.032)
12M Foreign Official Flows	-1.327***	(0.271)	-3.722***	(0.227)	-0.430	(0.399)	-2.215***	(0.274)
10Y Global Yield			0.925***	(0.072)			0.782***	(0.069)
3M Global Yield			-0.422***	(0.059)			-0.271***	(0.060)
Adj. R^2	0.910		0.948		0.915		0.946	
T	266		266		240		240	
ADF Statistic	-4.79***		-4.316***		-5.249***		-4.618***	

Regression estimates from Equation 3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Coefficients and standard errors for foreign official flows are estimated over the sum of coefficients on ΔFO_{t-l} where $l = 0, ...11$ over the last 12 months. Foreign official flows variable is scaled by U.S. GDP lagged 12 months. Regressions include an intercept term and linear time trend. Augmented Dickey-Fuller (ADF) test on regression residuals reject the null hypothesis that residuals are nonstationary. The sample periods under TIC flows and benchmark-consistent flows are January 1999 to February 2021 and January 1999 to December 2018, respectively. Benchmark-consistent flows data are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

Table A.10: 10-Year U.S. Yield Regressions with Sample Starting from 1990

	Dependent Variable: 10Y U.S. Yield							
	TIC Flows				Benchmark-Consistent Flows			
3M U.S. Yield	0.286***	(0.028)	0.282***	(0.024)	0.370***	(0.031)	0.363***	(0.026)
1Y GDP Forecast	0.011	(0.062)	0.059	(0.063)	0.353***	(0.084)	0.420***	(0.073)
10Y Inflation Forecast	0.488***	(0.166)	0.067	(0.165)	0.933***	(0.183)	0.533***	(0.192)
1Y Inflation Forecast	0.025	(0.058)	0.055	(0.054)	-0.087*	(0.049)	-0.082*	(0.043)
VIX	-0.013***	(0.004)	-0.008**	(0.004)	0.002	(0.005)	0.009**	(0.004)
Budget Surplus	-0.016	(0.032)	0.044*	(0.025)	-0.026	(0.035)	0.038	(0.030)
12M Foreign Official Flows	0.377**	(0.158)	-1.039***	(0.102)	0.153	(0.164)	-0.634***	(0.123)
10Y Global Yield			0.628***	(0.057)			0.603***	(0.051)
3M Global Yield			-0.228***	(0.031)			-0.204***	(0.032)
Adj. R^2	0.952		0.968		0.956		0.971	
T	374		374		348		348	
ADF Statistic	-5.323***		-4.562***		-6.034***		-5.427***	

Regression estimates from Equation 3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Coefficients and standard errors for foreign official flows are estimated over the sum of coefficients on ΔFO_{t-l} where $l = 0, ...11$ over the last 12 months. Foreign official flows variable is scaled by U.S. marketable debt lagged 12 months. Regressions include an intercept term and linear time trend. Augmented Dickey-Fuller (ADF) test on regression residuals reject the null hypothesis that residuals are nonstationary. The sample periods under TIC flows and benchmark-consistent flows are January 1990 to February 2021 and January 1990 to December 2018, respectively. Benchmark-consistent flows data are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

Table A.11: 10-Year U.S. Yield Regressions with Foreign Official UST Holdings

	Dependent Variable: 10Y U.S. Yield			
	Benchmark-Consistent Flows			
3M U.S. Yield	0.317***	(0.030)	0.295***	(0.021)
1Y GDP Forecast	0.314***	(0.103)	0.147*	(0.082)
10Y Inflation Forecast	0.176	(0.590)	-0.383	(0.362)
1Y Inflation Forecast	-0.019	(0.072)	-0.093***	(0.051)
VIX	0.000	(0.006)	-0.008	(0.005)
Budget Surplus	-0.088***	(0.030)	0.013	(0.020)
Foreign Official Holdings	-0.055***	(0.013)	-0.105***	(0.011)
10Y Global Yield			0.927***	(0.060)
3M Global Yield			-0.227***	(0.057)
Adj. R^2	0.915		0.960	
T	240		240	
ADF Statistic	-5.056***		-4.623***	

Regression estimates from Equation 3. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Foreign official holdings variable is a percentage of U.S. marketable debt lagged 12 months. Regressions include an intercept term and linear time trend. Augmented Dickey-Fuller (ADF) test on regression residuals reject the null hypothesis that residuals are nonstationary. The sample period under benchmark-consistent flows is January 1999 to December 2018. Benchmark-consistent flows data are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

Table A.12: Long-run Impact of Foreign Official Flows on 10-year U.S. Yields under Alternative Regression Specifications

Specification	TIC Flows Estimate	TIC Flows Std. Error	BC Flows Estimate	BC Flows Std. Error
Alternative global factor weighting				
GDP-weighted	-2.024***	(0.094)	-1.105***	(0.145)
Debt-weighted	-2.385***	(0.109)	-1.386***	(0.143)
CSA-weighted	-1.589***	(0.092)	-0.821***	(0.146)
PCA-weighted	-1.714***	(0.083)	-0.926***	(0.131)
Adjusting for 2-way yield feedback				
GDP-weighted, lagged	-1.942***	(0.100)	-1.037***	(0.148)
GDP-weighted, orthogonal	-0.834***	(0.095)	-0.478***	(0.127)
GDP-weighted, $y^{10Y}_{us,t} - \mathcal{Y}^{10Y}_{g,t}$	-2.287***	(0.099)	-1.448***	(0.143)
Additional conditioning variables				
CIP Deviations	-1.721***	(0.072)	-0.827***	(0.121)
Geopolitical Risk	-1.676***	(0.090)	-0.987***	(0.136)
Fed Forward Guidance	-2.027***	(0.086)	-1.134***	(0.140)
Fed Forward Guidance + LSAP	-2.778***	(0.140)	-2.413***	(0.097)
World ex. U.S. IP	-2.189***	(0.101)	-0.973***	(0.150)
Emerging Market Yields	-2.025***	(0.096)	-1.111***	(0.145)

Regression estimates from Equation 3 and estimates correspond to coefficients on '12M Foreign Official Flows' from Table 1 under TIC flows (left) and Benchmark Consistent (BC) flows of Bertaut and Tryon [2007] and Bertaut and Judson [2014] (right). The first group of estimates are based on long-term and short-term global yield factors constructed under alternative weight schemes. GDP refers to the benchmark GDP-weighted average global yield factors, hence the estimate is the same as that from Table 1. Debt refers to national public debt weighted average of international bond yields. CSA refers to the simple equal-weighted cross-section average. PCA refers to weights assigned via Principal Component Analysis. The second group of estimates adjust for potential reverse causality from U.S. yields to global non-U.S. yield factors. 'GDP-weighted, lagged' refers to GDP-weighted yield factors which enter the regression lagged one month. 'GDP-weighted, orthogonal' replaces the 10-year and 3-month factors in the regression with the residuals from linear regressions of the 10-year global yield factor on 10-year U.S. yields and the 3-month global yield factor on 3-month U.S. yields, respectively (Equation 6). 'GDP-weighted $y^{10Y}_{us,t} - \mathcal{Y}^{10Y}_{g,t}$' refers to a regression where the dependent variable is the U.S.-global 10-year yield differential instead of the 10-year yield alone. The third group of estimates control for additional global covariates. U.S. Treasury covered interest rate parity (CIP) deviations, or the cross-currency basis by augmenting the regression with the average 3-month and 10-year Treasury CIP deviations across G10 countries, taken from Du and Schreger [2016] and Du et al. [2018]. Geopolitical risk is measured using the news-based index of Caldara and Iacoviello [2018]. Federal reserve forward guidance and Large scale asset purchase (LSAP) shocks are measured as in Swanson [2021] and two covariates enter the regression for each variable: the forward guidance and LSAP shocks, and a cumulative sum of the forward guidance shocks and LSAP shocks, respectively. World ex. U.S. IP is a measure of year-on-year global industrial production growth excluding the United States. Emerging market yields refer to monthly changes in the J.P. Morgan EMBI sovereign spread index to capture shocks in emerging markets. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance, respectively. Coefficients and standard errors for foreign official flows are estimated over the sum of coefficients on $\Delta FO_{t=l}$ where $l = 0, ..., 11$ over the last 12 months. Foreign official flows variable is scaled by GDP lagged 12 months. Regressions include an intercept term and linear time trend. The sample period is January 1999 to February 2021 for TIC flows, and January 1999 to December 2018 for BC flows.

A6 VAR Specification and Identification

A6.1 Discussion on VAR specification

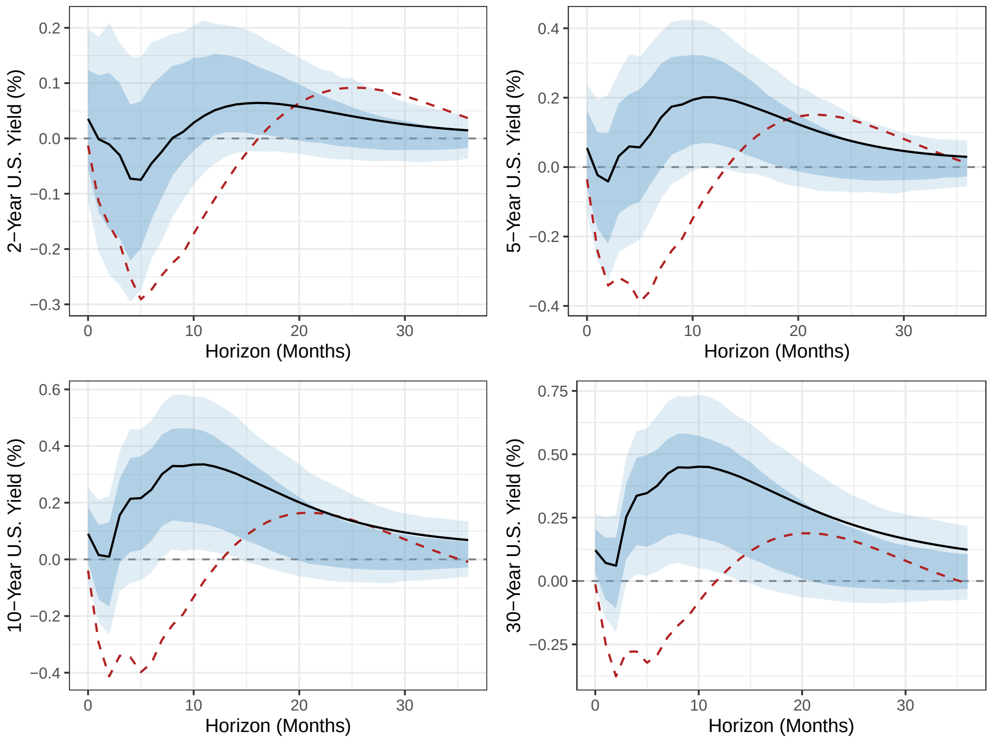
An important reason to estimate individual yields in the unrestricted VAR system due to its simplicity, transparency, and relatively few assumptions. Alternatively, one can adopt more sophisticated approaches to model the U.S. term structure such as affine or equilibrium modeling for the yield curve which impose no-arbitrage restrictions across maturities [Kohn, 2016; Kaminska and Zinna, 2020; Wolcott, 2020]. A key benefit of these term structure models is that they are theoretically grounded, often useful for recovering measures of market-based inflation expectations or term premia from the yield curve. Yield curve decomposition, however, is not the purpose of this study. Moreover, estimation of no-arbitrage term structure models remains a significant challenge, as different estimation procedures can lead to substantial differences in parameter estimates (and therefore widely different overall results) due to the highly non-linear nature of the likelihood surfaces [Hamilton and Wu, 2012]. To impose no-arbitrage across the yield curve, further modeling choices are required such as which tenors are priced with and without error and whether or not two-way feedback is allowed between yields and macro variables. It is also not clear whether these approaches offer material advantages in yield curve modeling over approaches which do not impose similar structural restrictions [Carriero and Giacomini, 2011; Favero et al., 2012], especially in the large, highly liquid market for USTs [Coroneo et al., 2011]. Therefore to prioritize simplicity and robustness, an unrestricted yield curve VAR is considered, the simplest dynamic extension of regression specification (3).

A6.2 VAR with zero restrictions

As an alternative to identification through heteroskedasticity, the VAR specification from Section 4 is identified with short-run zero restrictions. First, Figure A.8 IRFs order foreign official flows first in the VAR. Second, to further guard against the threat of simultaneity, foreign official are flows ordered last rather than first IRFs reported in Figure A.9. Global factors are also controlled for in all specifications.

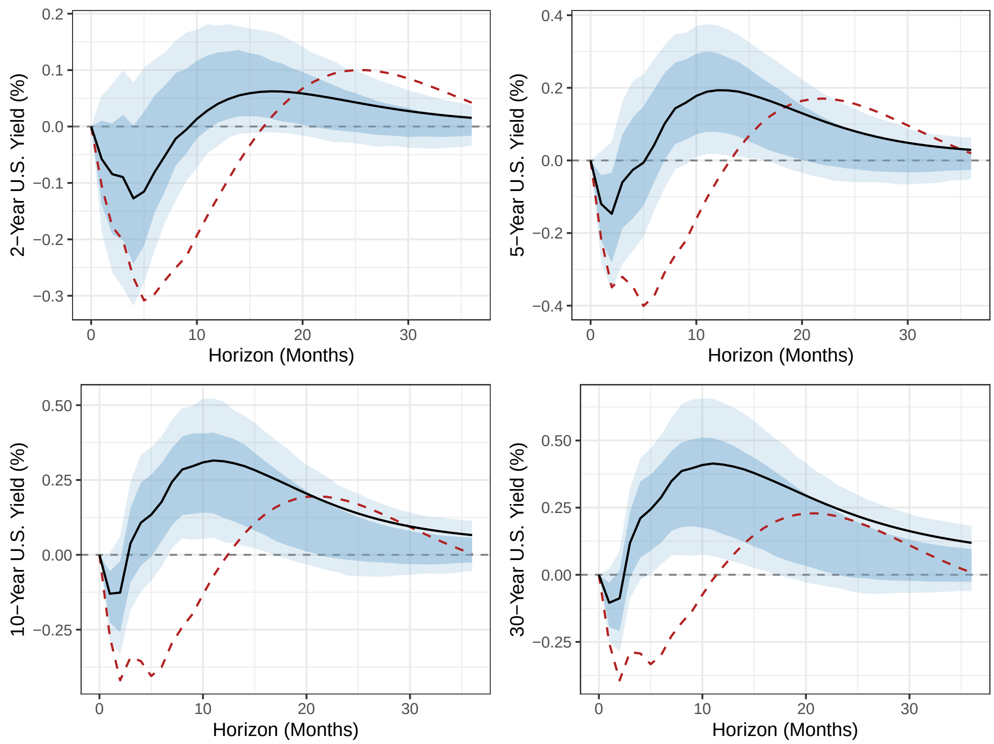
Figure A.10 plots the IRFs following a sale of USTs the size of 1% of marketable debt by foreign private investors. Specifically, the VAR from specification (4) is extended to include an equation for foreign private UST flows. To deal with the elastic nature of private UST demand, foreign private flows are ordered last in the VAR while official flows remain ordered first. Foreign private sales induces positive but smaller responses in U.S. yields, consistent with their relatively elastic demand compared to foreign officials.

Figure A.8: Impulse Response of U.S. Yields over the Federal Funds Rate to a Foreign Official UST Sale (1% of Debt) Ordered First in the SVAR



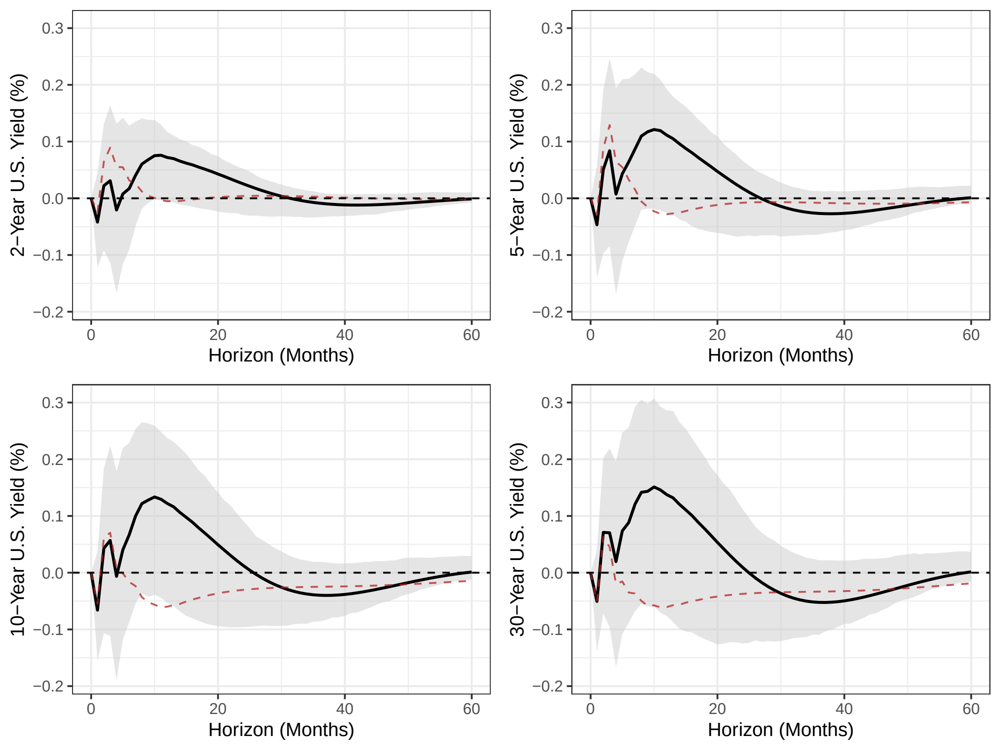
Impulse responses from a VAR(4) specified as in Equation 4 using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. Light and dark shaded regions refer to 90% and 68% bootstrapped confidence interval based on 1,000 bootstrapped samples. Dashed line reflects impulse response from VAR(4) that excludes global yield factors from the set of contemporaneous controls. Foreign official UST flows are ordered first in a Cholesky decomposition identifying structural shocks.

Figure A.9: Impulse Response of U.S. Yields over the Federal Funds Rate to a Foreign Official UST Sale (1% of Debt) Ordered Last in the SVAR



Impulse responses from a VAR(4) specified as in Equation 4 using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. Light and dark shaded regions refer to 90% and 68% bootstrapped confidence interval based on 1,000 bootstrapped samples. Foreign official UST flows are ordered last in a Cholesky decomposition identifying structural shocks.

Figure A.10: Impulse Response of U.S. Yields over the Federal Funds Rate to a Foreign Private UST Sale (1% of Debt) Ordered Last in the SVAR



Impulse responses from a VAR(4) specified as in Equation 4 using benchmark-consistent foreign official flows, augmented with foreign private flows as an additional endogenous variable following Bertaut and Tryon [2007] and Bertaut and Judson [2014]. Both flow measures are scaled by U.S. marketable debt lagged 12 months. Foreign official flows ordered first in the VAR and foreign private flows are ordered last in the VAR. Shaded region refers to 90% bootstrapped confidence interval based on 1,000 bootstrapped samples.

A7 Spillovers from China's Exchange Rate Policy

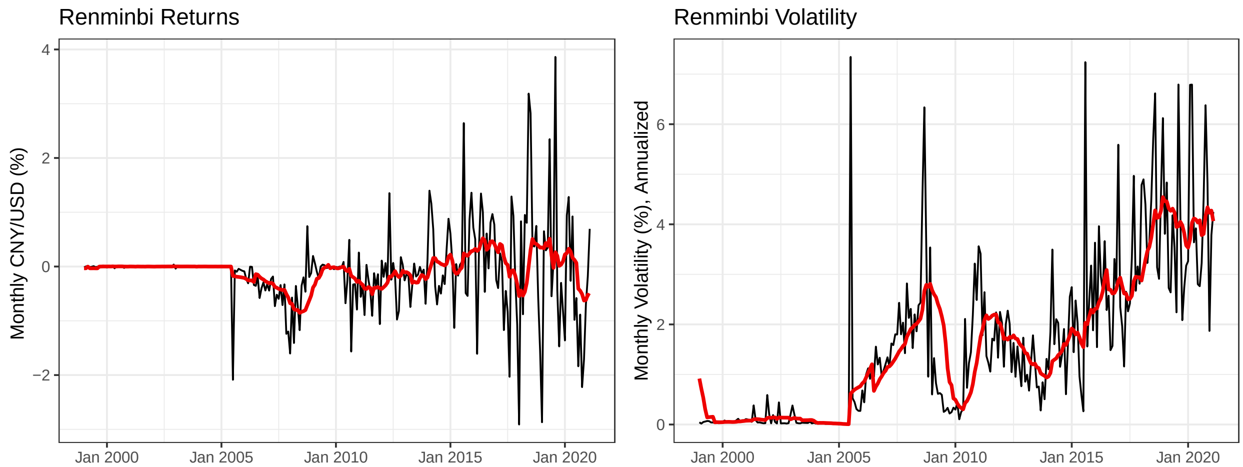
For the estimates reported thus far to be considered causal, two important conditions must be satisfied, assuming minimal measurement error: first that official demand for UST is not affected by the price of U.S. Treasury securities (the simultaneity problem) and second, that global factors properly condition on unobserved factors which jointly affect UST demand and U.S. Treasury yields (the omitted variables problem). Global yield factors aims to address the latter condition while thus far the former condition has been addressed by implicitly assuming foreign official UST demand is inelastic. This section relaxes the former assumption explicitly. The literature has previously considered event studies or instrumental variables (IVs) to address the simultaneity problem in case the assumption of inelastic UST demand is violated [Bernanke et al., 2004; Beltran et al., 2013; Martin, 2014b; Fang and Liu, 2019]. In view of that consideration, this section extends the VAR analysis and deals with the possibility of simultaneity by exploiting *lagged* monthly realized volatility of the Chinese Renminbi - U.S. Dollar exchange rate (CNY), while continuing to condition on global yield factors. That way, the updated empirical approach tries to achieve cleaner identification via a two-pronged strategy that addresses simultaneity bias and omitted variable bias at the same time. While interpreting these results, note the difficulty of identifying experiment-quality exogenous variation at the international macroeconomic scale, and as such, these results should be interpreted as a small step toward the ideal causal setting rather than a perfect solution.[A10]

China can be considered a granular investor in the U.S. Treasury market [Gabaix and Koijen, 2020], amassing a large percentage of all foreign exchange reserves accumulated by official institutions, dwarfing most countries.[A11] The country's mercantilist policies amid

[A10] The presence of spillovers and general equilibrium effects make most if not all identification strategies in empirical macroeconomic settings at the global level imperfect because of the difficulty of satisfying the 'stable unit treatment value assumption' (SUTVA).

[A11] In this context, a 'granular' economic agent is one that is not atomistic, and therefore does not take market prices as given. Rather, a granular agent is large or influential enough such that its decisions can affect or determine the market price.

Figure A.11: China-U.S. Exchange Rate Monthly Returns and Volatility



Smoothed line is the 12-month trailing average. Left panel shows CNY/USD monthly returns where negative values correspond to Chinese Renminbi appreciation relative to the U.S. Dollar. Right panel shows monthly CNY/USD realized volatility, annualized. Monthly returns are the sum of daily log CNY/USD returns, and realized volatility is computed as the standard deviation of daily CNY/USD returns within each month (Equation A.7).

export-led growth led to substantial management of its exchange rate, intentionally limiting currency appreciation to maintain international competitiveness. As a result, motives to smooth out exchange rate volatility crucially shaped China's demand for international reserves during the 2000s, much of which ended up in U.S. Treasury securities [Ito et al., 2015; Das, 2019].[A12] Consistent with the 'global savings glut' hypothesis, this in turn put downward pressure on U.S. interest rates [Obstfeld and Rogoff, 2009; Eugeni, 2015]. Hence, details on Chinese Renminbi exchange rate regimes may be a valid proxy for foreign official UST demand that deals with the simultaneity problem. In August 2015, China shifted policy toward a more flexible Renminbi regime to be determined by market forces [McCauley and Shu, 2019]. This came alongside China's intent to have the Renminbi included in the International Monetary Fund's (IMF) SDR basket, which was announced October 1, 2016. These regime shifts imply that Renminbi volatility varied remarkably over the sample period

[A12] As of 2018, China (both official and private sector) owned a total of $1.18 trillion in UST securities, about 19% of all Treasury securities held abroad – the largest foreign holder of USTs. As of June 2020 China held 17.3% of all long-term Treasury securities held abroad, now second to Japan. China holds roughly 5% of all Treasury securities outstanding.

under consideration, from a *de facto* hard Dollar peg in the early 2000s and again during the 2008 GFC (despite a *de jure* manage float policy in 2005), to a managed float by 2016. We specifically consider the volatility of the Renminbi exchange rate an indicator of the country's *de facto* time-varying exchange rate regime linked to UST demand. Higher volatility from a flexible exchange rate policy meant a decreasing need to manage the exchange rate through accumulation of international reserves or active intervention. As supplementary evidence, Figure A.15 documents that greater Renminbi volatility (increased FX flexibility) leads to declining FX reserves held by China. Based on this rationale, the relationship between foreign official UST demand and Renminbi returns and volatility, respectively, can be summarized as follows:

1. *Exchange rate smoothing implies that larger CNY depreciation (appreciation) is followed by larger UST sales (purchases).*
2. *Low CNY volatility (e.g. U.S. Dollar peg) and high CNY volatility (e.g. floating) regimes imply relatively higher and lower demand for USTs, respectively.*

Lagged measures of the Renminbi currency's distribution have not been previously considered as proxies to address simultaneity as the literature has mainly relied on Japanese FX interventions, oil supply shocks, China trade and investment flows, or changes in China's international reserves holdings as instruments. The latter two related to China are the closest analog to using lagged Chinese exchange rate fluctuations. However, tests indicate that both China's trade balance/investment flow and international reserves data are weak instruments for official UST flows [Beltran et al., 2013; Fang and Liu, 2019] which could be for multiple reasons. Perhaps fluctuations in these related series result from large general equilibrium effects and spillovers, dominating any exogenous variation linked to official flows. Or the slow-adjusting, smoother nature of real macroeconomic data relative to official flows and market-based U.S. yields create differences in measured variation that weaken the statistical relationship between China trade or financial flows with foreign official UST flows. Another possibility for the weak instruments problem is measurement error due to mis-reporting of official Chinese data [Nakamura et al., 2016; Fernald et al., 2021].

Because its value updates at higher frequencies and the transparent nature of financial market data, exploiting information in Renminbi variation partly remediates these issues and therefore could lend to greater instrument relevance, while still exploiting the dominant role of China in the UST market. That said, it's worth noting that while lagged Renminbi volatility may be useful for addressing simultaneity, the endogenous nature of currency movements mean that it does not represent truly exogenous variation just like the other instruments used in the literature. Moreover, all such 'instruments' may be influenced by global confounding factors. Therefore, proper identification still requires conditioning on global yield factors to account for the global component of U.S. yields and potential cyclicality of foreign official UST demand.

Using daily data on the Renminbi-U.S. exchange rate (CNY/USD), we compute daily currency log returns and then aggregate to monthly returns and realized volatility:

$$\Delta cny_t = \sum_{d \in t} \ln \frac{CNY_d(t)}{CNY_{d-1}(t)}, \qquad \sigma_t^{cny} = \left(\frac{1}{D(t)-1} \sum_{d \in t} [\ln \frac{CNY_d(t)}{CNY_{d-1}(t)} - \Delta \widetilde{cny}(t)]^2 \right)^{1/2}, \tag{A.7}$$

where Δcny_t is the monthly log CNY/USD return computed as the sum of daily returns within each month, where $CNY_d(t)$ refers to the CNY/USD exchange rate on day d in month t. The exchange rate is in terms of Renminbi per 1 U.S. Dollar, hence negative values of log returns correspond to CNY appreciation relative to USD. The realized month t volatility of the exchange rate is given by σ_t^{cny}, which is computed as the sample standard deviation of daily log CNY/USD returns within month t, where $\Delta \widetilde{cny}(t)$ refers to the average of daily Renminbi returns in month t, or Δcny_t divided by the number of observations in month t. These series are plotted in Figure A.11 along with their 12-month moving averages. The correlation between Δcny_t and σ_t^{cny} is statistically indistinguishable from zero.

To test the relevance of Chinese exchange rate returns and volatility as predictors of foreign official UST flows, we provide regression estimates in Table A.13 from regressing foreign official flows on an intercept, lagged CNY/USD returns, and lagged volatility:

$$\Delta FO_{t-l} = \delta_0 + \delta_1 \Delta cny_{t-l-1} + \delta_2 \sigma_{t-l-1}^{cny} + e_{t-l}, \qquad \text{for } l = 0, ..., 11. \tag{A.8}$$

Both 1-month lagged Renminbi measures as well as the lagged 12-month averages are considered. As a rule of thumb in the IV literature, Staiger and Stock [1997], suggest an F-test statistic of 10 or greater to reduce the risk of weak instruments.[A13] The reported statistics of 28.09 for lagged CNY returns and volatility, and 60.9 for the lagged 12-month trailing average return and volatility both are satisfactorily large, and the results hold under both TIC flows or benchmark-consistent flows. Lagged Chinese Renminbi variables explain a substantial portion of variation in foreign official flows, ranging from nearly 10% to 31% across specifications.

Table A.13: Chinese Renminbi Returns and Volatility and Foreign Official UST Flows

	Dep. Var: TIC Flows				*Dep. Var:* Benchmark-Consistent Flows			
Lagged CNY Returns	-4.992***	(1.741)			-6.850***	(2.286)		
Lagged CNY Volatility	-4.065***	(0.718)			-3.271***	(0.981)		
Lagged CNY Returns (12M Avg.)			-23.813***	(4.220)			-28.591***	(5.173)
Lagged CNY Volatility (12M Avg.)			-6.001***	(0.764)			-6.868***	(1.232)
Adjusted R^2	0.17		0.311		0.087		0.211	
T	266		266		240		240	
F Statistic	28.09		60.90		12.37		33.09	

Standard errors are adjusted for heteroscedasticity and autocorrelation. *,**,*** correspond to 10%, 5% and 1% significance, respectively. Regressions include an intercept term. The sample periods under TIC flows and benchmark-consistent flows are January 1999 to February 2021 and January 1999 to December 2018, respectively. Benchmark-consistent flows are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

The estimates reported in Table A.13 are also consistent with China's exchange rate smoothing and mercantilist motives. A negative estimate on CNY returns implies that Renminbi appreciation was followed by foreign official purchases of U.S. Treasuries, which would entail selling CNY for USD to prevent further Renminbi appreciation. Meanwhile

[A13] This was generalized to the case of multiple instruments by Stock and Yogo [2005].

a negative estimate on CNY volatility suggests that a more flexible exchange rate regime necessitates less exchange rate management – leading to weaker official UST demand.[A14]

A7.1 Incorporating Renminbi volatility in the structural VAR

We focus on CNY volatility (σ_t^{cny}) rather than monthly returns (Δcny_t) for the structural VAR analysis. While both returns and volatility fluctuations are endogenous, realized volatility has a more structural interpretation related to China's exchange rate regime and is presumably less sensitive to endogeneity concerns that currency returns and interest rates are subject to. For robustness, Figure A.14 reports consistent evidence using CNY returns. It is also worth reiterating that CNY realized volatility is uncorrelated with CNY exchange rate returns.

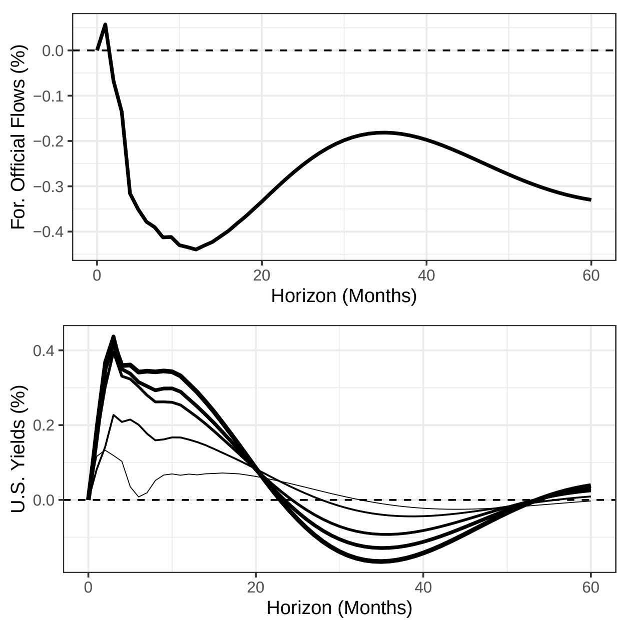
Recent advances of Mertens and Ravn [2013] and others exploit exogenous variation in IVs within a VAR framework. However, for this proxy-VAR approach to be applied appropriately, the source of external variation must be exogenous. Because realized CNY volatility is still likely to be endogenous, a second-best approach would have CNY volatility enter the VAR described in (4) as an endogenous variable, ordered last. CNY volatility is ordered last to address simultaneity concerns, imposing the restriction that foreign official UST demand (and the impact on U.S. yields) can only respond to shifts in CNY volatility with a lag. Contemporaneous domestic and global factors enter as conditioning variables.

The IRFs to a CNY volatility shock will induce a joint, lagged response of foreign official UST flows and U.S. yields. Intuitively, a CNY volatility shock can be viewed as a sudden shift in the level of exchange rate volatility, or a Renminbi exchange rate regime shift. The underlying assumption for identification is that *lagged* CNY volatility shocks can only impact U.S. yields through their effect on foreign official UST demand. If markets are informationally efficient, any information-driven changes in CNY volatility that also affects

[A14]Under a more flexible exchange rate, the central bank's demand for USTs may decrease because a lower stock of international reserves is required than that under a peg. In addition, as the exchange rate becomes market-determined, interventions are less frequently required.

U.S. interest rates should occur contemporaneously. Therefore, information in lagged CNY exchange rate fluctuations should be embedded in U.S. yields from the same period, but should not be able to forecast U.S. interest rates in future periods. Given the size and depth of the U.S. Treasury market, this may be a reasonable assumption. However, if changes in CNY volatility prompt officials to buy or sell USTs, this operation may give rise to a lead-lag relationship between the Renminbi and official UST demand or U.S. yields as shown. Moreover, the response of flows and yields need not be synchronous. For example, flows may be affected for several periods following a shock while the U.S. yield response might peak quickly. This merely suggests that asset prices incorporate changing future flows in a forward-looking manner. Figure A.12 shows suggestive evidence supporting the "official flows channel" through which CNY volatility can impact U.S. yields: the joint responses of foreign official flows and U.S. yields to a CNY shock appear to mirror each other.

Figure A.12: Cumulative Response of Foreign Official Flows and U.S. Yields to a CNY Volatility Shock of 5%



Impulse responses to a 5% CNY volatility shock from a VAR(4) specified as in Equation 4 which includes realized CNY volatility as an endogenous variable computed as in Equation A.7 and using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. CNY volatility is ordered last in the VAR. Upper panel: Response of foreign official UST flows is cumulative. Lower panel ordered from thin to thick: 3M, 2Y, 5Y, 10Y, 30Y U.S. yield responses.

Figure A.13 reports IRFs from a VAR following a CNY volatility shock of 5% (annualized volatility). Given the standard deviation of monthly CNY realized volatility is 1.5%, this can be considered a quantitatively large shock. Figure A.11 shows that over the last two decades as the Renminbi transitioned from a pegged to a floating regime, CNY realized volatility increased from about 0% to 4%, exceeding 6% on certain occasions. Faced with an increase in Renminbi flexibility of this size, foreign official UST flows (presumably linked to China) are positive for one month, followed by persistent UST selling by foreign officials in 10 out of the next 11 months. At its trough, foreign officials liquidate USTs in the cumulative amount of 0.44 percent of marketable debt outstanding following a CNY volatility shock of 5% (Figure A.12). There is also a significant rise in U.S. yields following a CNY volatility shock, with 2-year yields rising just about 20 basis points but longer-term yields rising 40 basis points. The impact on yields is realized relatively quickly, peaking within 4 months of the CNY shock. Assuming the impact on U.S. yields is completely driven by changes in foreign official UST flows linked to China, a back-of-the-envelope calculation suggests that foreign official UST sales of 1 percent of marketable debt is associated with a rise in 10-year yields of 90.9 basis points[A15], or a \$100 billion flow is associated with a 50 basis point yield change when China's exchange rate policy is the source of the shock (assuming \$18 trillion in marketable debt). This is roughly double the size of the estimated 27.5 basis point impact Section 4 which does not explicitly account for simultaneity bias or single out China-driven UST demand.

The dashed red IRFs report the response of U.S. yields when we do not control for global yield factors. The estimated impact on U.S. yields is much smaller, often statistically indistinguishable from zero when global factors go unaccounted for. Not shown for brevity, but using the 12-month rolling average of CNY realized volatility in the VAR to smooth out idiosyncratic monthly volatility shocks which may be driven by one-off devaluations leads to similarly shaped and larger IRFs in response to the same 5% volatility shock. In Section

[A15] 40 basis points $\times \frac{1}{0.44}$.

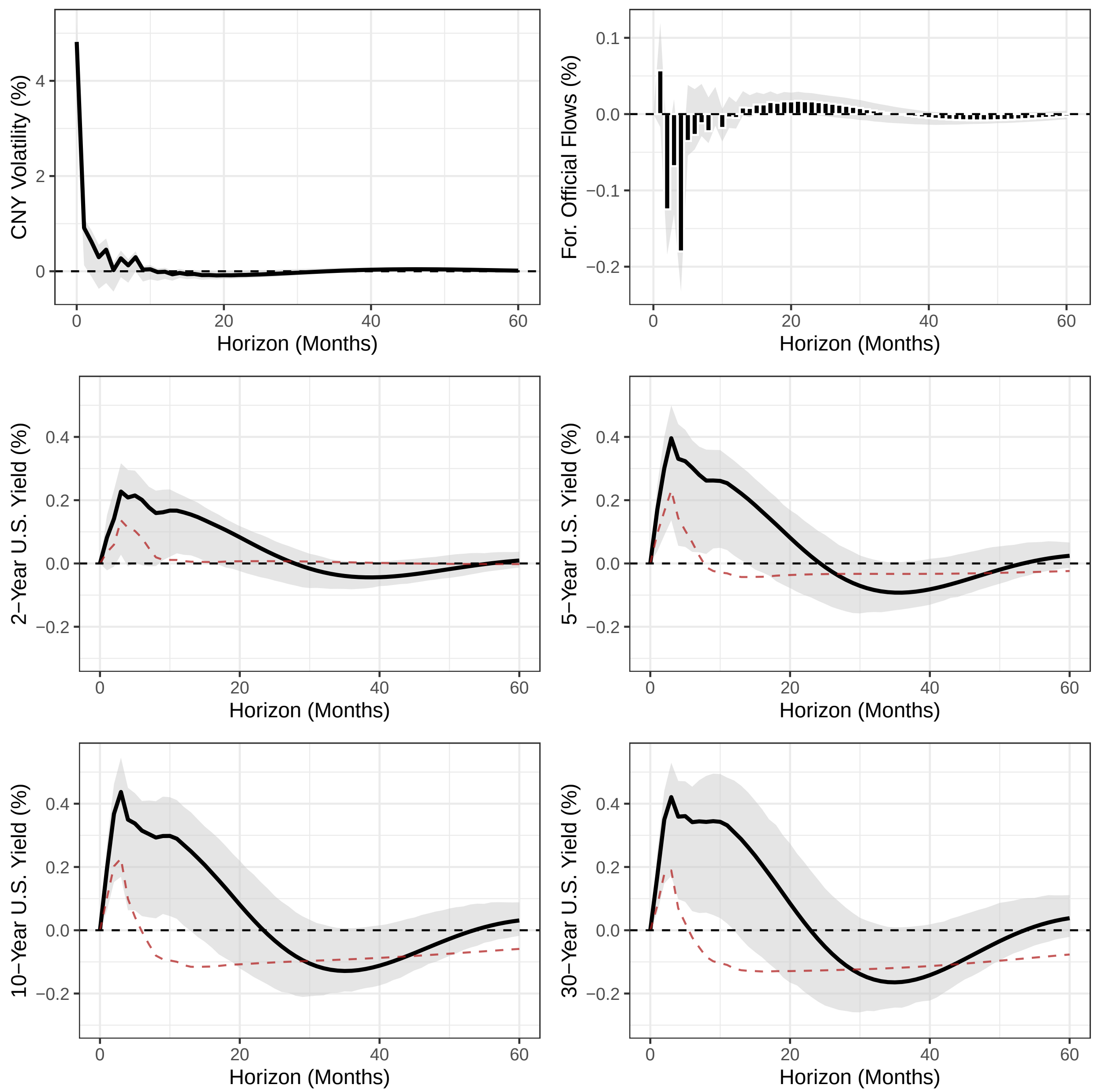
A7.3 of the Appendix, we test for potential structural breaks starting in 2015 within the relationship between Renminbi fluctuations and foreign official UST flows. 2015 marks the starting period of major Chinese Renminbi reform and internationalization. Consistent with this, regression estimates find significant evidence of a structural break in the Renminbi-official UST flows relationship, with it strengthening significantly since 2015. However, Figure A.16 shows that the VAR results following a CNY volatility shock remains robust when estimated over a pre-2015 sample, implying that the post-2015 period is not driving the overall results.

A7.2 Supporting evidence on the role of China's exchange rate policy

As supplementary evidence linking China's exchange rate movements to their official demand for USTs (and related impact on U.S. yields), we provide two additional statistics. First, Renminbi depreciations (appreciations) should be associated with lower (higher) demand for USTs from China, if currency fluctuations do indeed portend meaningful intervention. Replacing realized CNY volatility in (A.7) with monthly CNY returns, Figure A.14 plots the impulse response of foreign official UST flows (left) and the U.S. yield curve (right) following a 1% CNY depreciation shock. The shock is ordered so there is zero contemporaneous effect. Consistent with this hypothesis, CNY depreciation leads to foreign official UST sales and an increase in U.S. yields after an initial, short-lived drop similar to results found in Figure A.8. Longer maturity yields eventually head lower after the increase, while shorter maturity yields remain elevated, thereby a CNY depreciation leads to a flatting of the U.S. yield curve over time.

The second statistic supports the view that China's exchange rate regime can be reasonably approximated by the volatility of the exchange rate. In a simple two-variable VAR(4) with Renminbi realized volatility ordered second, Figure A.15 reports that increased Renminbi volatility is associated with future declines in Chinas foreign exchange reserves (ex-

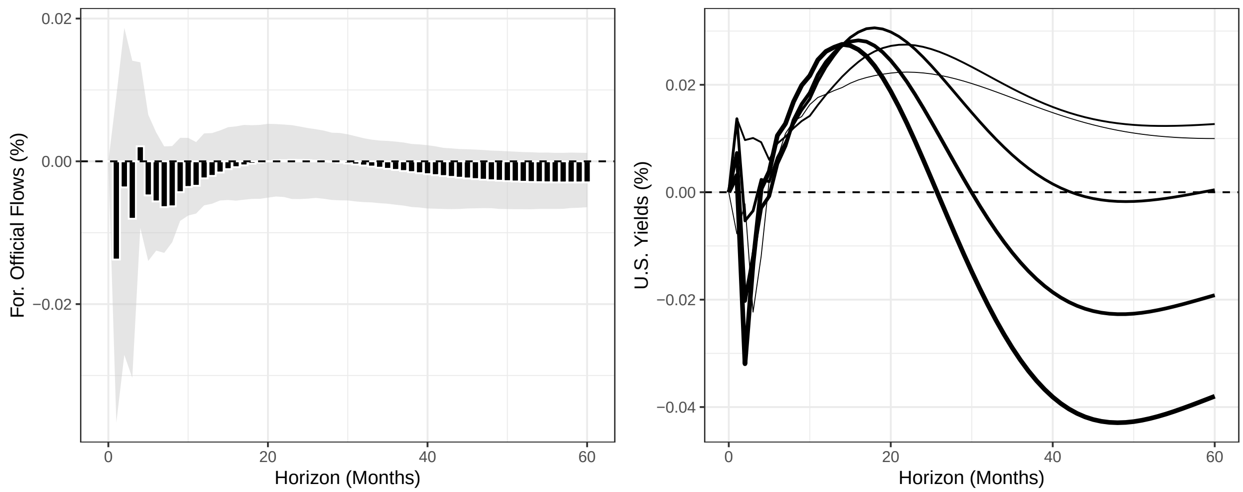
Figure A.13: Impulse Responses to a CNY Volatility Shock of 5% (annualized)



Impulse responses from a VAR(4) specified as in Equation 4 which includes realized CNY volatility as an endogenous variable computed as in Equation A.7 and using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. Shaded region refers to 90% bootstrapped confidence interval based on 1,000 bootstrapped samples. Dashed line reflects impulse response from VAR(4) that excludes global yield factors from the set of contemporaneous controls. CNY volatility is ordered last in the VAR.

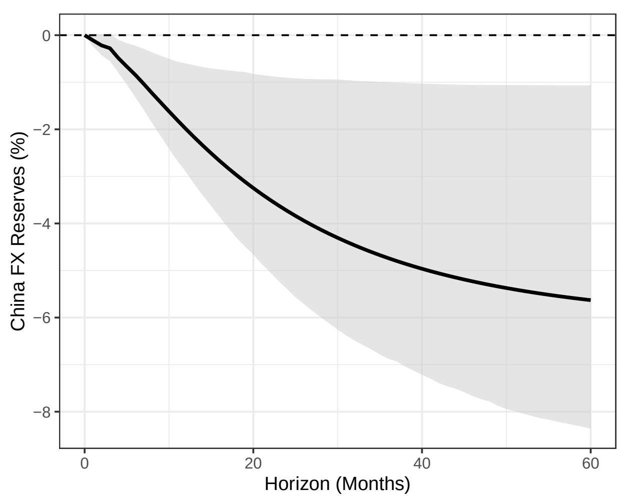
cluding gold). This pattern ties together the connection between China's Renminbi volatility and total foreign official UST flows, which operates through China's managing of their FX

Figure A.14: Response of Foreign Official Flows and U.S. Yields to a CNY Depreciation Shock of 1%



Impulse responses to a 1% CNY depreciation shock from a VAR(4) specified as in Equation 4 augmented with CNY returns as an endogenous variable computed as in Equation A.7 and using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. CNY returns ordered last in the VAR. Left panel: Response of foreign official UST flows, with shaded region referring to 90% bootstrapped confidence interval based on 1,000 bootstrapped samples. Right panel ordered from thin to thick: 3M, 2Y, 5Y, 10Y, 30Y U.S. yield responses.

Figure A.15: Response of China's FX Reserves to a CNY Volatility Shock (5% annualized)



Impulse responses to a 5% CNY volatility shock from a two-variable VAR(4) of CNY realized volatility and China's FX reserves excluding gold. CNY realized volatility is computed as in A.7 and ordered second in the VAR. Shaded region refers to 90% bootstrapped confidence interval based on 1,000 bootstrapped samples.

reserves. While breaking down China's FX reserves by currency composition and security type poses challenges due to data limitations, a large portion of their total reserves tend to be composed of USTs.

A7.3 Regime Shifts in The Renminbi-Official UST Flows Relationship

As noted in Section A7, China has undergone dramatic changes in its exchange rate policy over the past two to three decades. One important change occurred in 2015 as China attempted to globalize its currency and allow a greater degree of flexibility in the Renminbi. At the same time, the global share of international reserves held in U.S. Dollars has steadily declined from 66% in 2015 to about 61% in 2020 while the Renminbi entered the basket of global reserve currencies. These macro shifts in exchange rate policy suggest possible structural breaks in the relationship between Renminbi fluctuations and foreign official UST demand. While including the level of CNY volatility in these regressions aims to capture such regime shifts, it may not be sufficient. To test the presence of structural breaks beginning 2015, the regression from (A.8) and Table A.13 is augmented with regime indicators:

$$\Delta FO_t = \delta_0 + [\delta_1 + \delta_{1b}\mathbf{1}_{t\geq \text{Jan 2015}}]\Delta cny_{t-1} + [\delta_2 + \delta_{2b}\mathbf{1}_{t\geq \text{Jan 2015}}]\sigma^{cny}_{t-1} + e_t, \qquad (A.9)$$

where the indicator $\mathbf{1}$ receives a value of one in periods on or after January 2015, and zero otherwise. We consider both one-month lagged Renminbi returns and volatility along with lagged 12-month averages.

Table A.14 reports regression results under the augmented first-stage regressions allowing for a regime change in 2015. The estimates suggest a significant change in the relationship between Renminbi fluctuations and official flows during the post-2015 period. During this period specifically, both lagged CNY returns and volatility become increasingly negatively associated with one-month ahead official UST flows, and explanatory power after allowing for the regime change also rises substantially compared to the results reported in Table

Table A.14: Instrument Relevance Regressions: Allowing for a Regime Shift in 2015

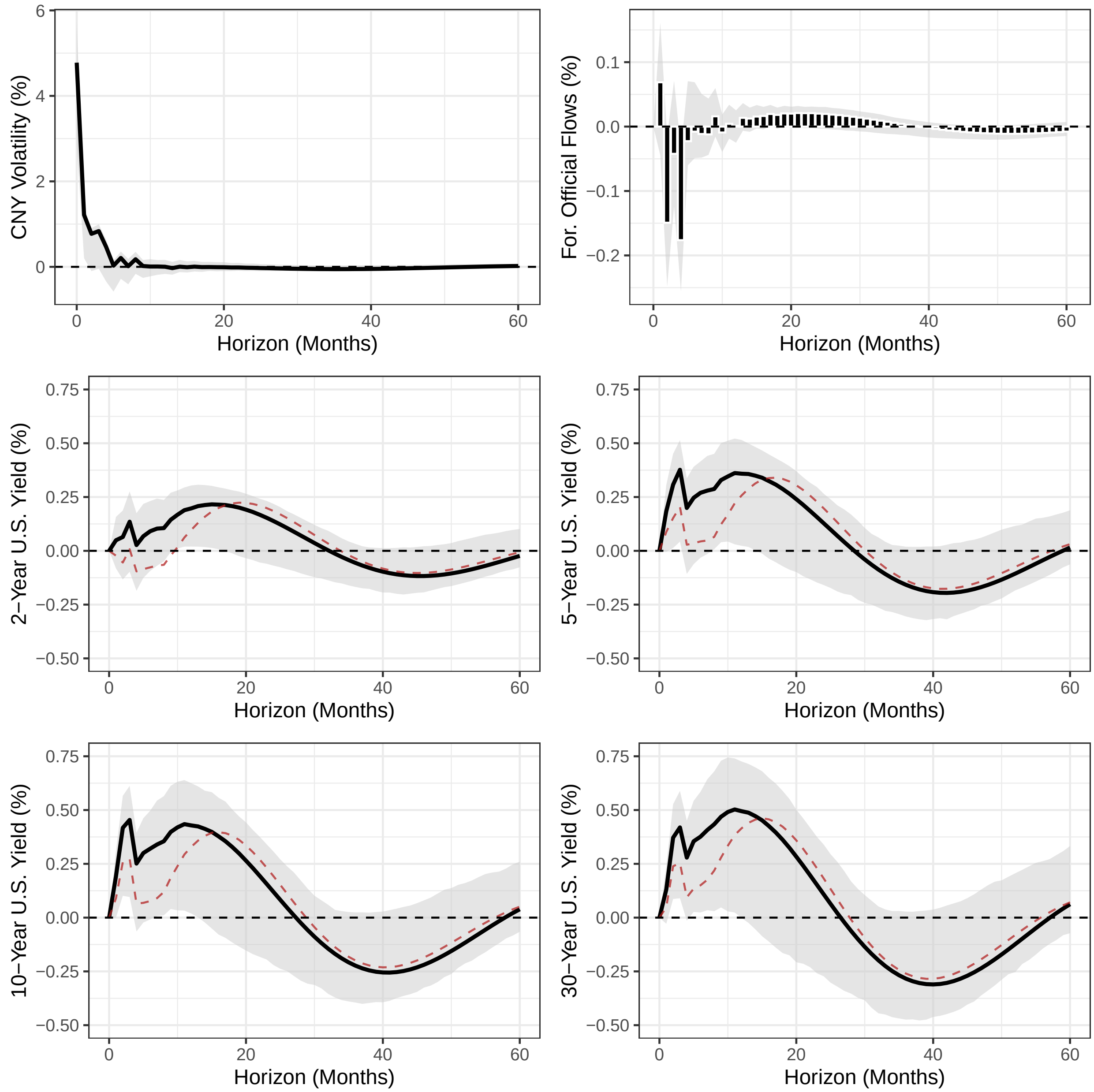
	Dep. Var: TIC Flows				*Dep. Var:* Benchmark-Consistent Flows			
Lagged CNY Returns	0.054	(3.043)			-0.647	(3.428)		
Lagged CNY Returns × Post-2015	-3.475	(3.376)			-2.177	(4.787)		
Lagged CNY Volatility	2.234**	(1.096)			2.681	(1.665)		
Lagged CNY Volatility × Post-2015	-7.486***	(1.076)			-8.799***	(1.805)		
Lagged CNY Returns (12M Avg.)			7.311	(7.691)			9.897	(10.833)
Lagged CNY Returns (12M Avg.) × Post-2015			-30.556***	(10.541)			-30.099*	(15.783)
Lagged CNY Volatility (12M Avg.)			3.43*	(2.040)			5.006*	(2.825)
Lagged CNY Volatility (12M Avg.) × Post-2015			-9.378***	(2.022)			-13.266***	(2.910)
Adjusted R^2	0.31		0.37		0.21		0.30	
T	266		266		240		240	
F Statistic	29.45		39.65		16.48		26.50	

Standard errors are adjusted for heteroscedasticity and autocorrelation. *,**,*** correspond to 10%, 5% and 1% significance, respectively. Regressions include an intercept term. The sample periods under TIC flows and benchmark-consistent flows are January 1999 to February 2021 and January 1999 to December 2018, respectively. Benchmark-consistent flows are based on Bertaut and Tryon [2007] and Bertaut and Judson [2014].

A.13 with no regime change. That is, Renminbi appreciation and lower Renminbi realized volatility are both associated with relatively larger foreign official UST purchases in the post-2015 period than before.

Figure A.16 reports IRFs from a structural VAR augmented with Renminbi realized volatility ordered last (as in Section A7). The sample period, however, excludes January 2015 onward to test whether the IRFs are sensitive to the post-2015 period under the Renminbi regime change. It is clear that while 2015-onward represents an important shift in Chinese exchange rate policy, the effects relating Renminbi volatility to foreign official UST demand and resulting U.S. Treasury yields were present and significant prior to 2015.

Figure A.16: Impulse Responses to a CNY Volatility Shock (5% annualized) using Pre-2015 Sample



Impulse responses from a VAR(4) specified as in Equation 4 which includes realized CNY volatility as an endogenous variable computed as in Equation A.7 and using benchmark-consistent foreign official flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] scaled by U.S. marketable debt lagged 12 months. Shaded region refers to 90% bootstrapped confidence interval based on 1,000 bootstrapped samples. Dashed line reflects impulse response from VAR(4) that excludes global yield factors from the set of contemporaneous controls. CNY volatility is ordered last in the VAR. CNY volatility is ordered last in identifying structural shocks. Data from January 1999 through December 2014, prior to the 2015 Renminbi reforms.