Dollar Reserves and U.S. Yields: Identifying the Price Impact of Official Flows*

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Abstract

This paper shows that the impact of foreign official demand for USTs on U.S. yields is much larger when critical sources of endogeneity are addressed. We exploit changes in the volatility of foreign official UST flows before and after the 2008 Global Financial Crisis to identify a VAR of U.S. yields via heteroskedasticity. A foreign official flow shock of $100B moves 5-year, 10-year, and 30-year yields by more than 100 basis points on impact, converging to the range of estimates previously reported in the literature within 6 months.

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

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1 Introduction

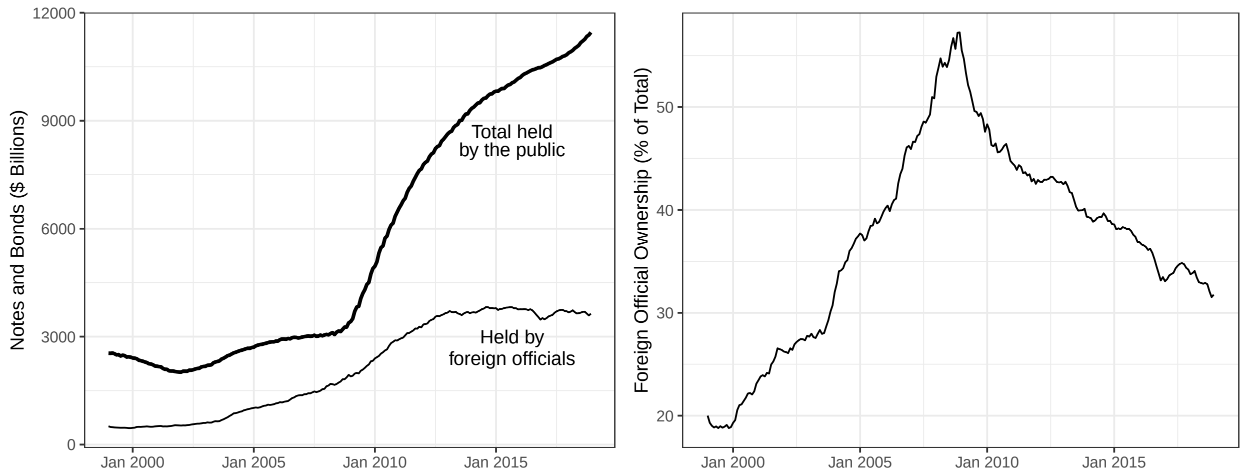
The 2008 Global Financial Crisis, COVID-19 pandemic, and resulting issuance of central bank swap lines revealed the financial stability risks of concentrated ownership of U.S. Treasuries (USTs) by foreign officials (Figure 1).[1] These episodes not only confirm the inelasticity of foreign official UST demand [Alfaro et al., 2014; Tabova and Warnock, 2021] but also point toward its highly endogenous nature. Foreign official demand is pro-cyclical, determined by economic forces that also affect Treasury yields, giving rise to the presence of confounding factors. Demand for Treasuries affect Treasury yields but yields also shape demand, especially demand from price-elastic investors. This results in the simultaneous determination of Treasury demand and Treasury prices. These sources of endogeneity – confounding (or omitted) variables and simultaneity – make consistently estimating the price impact of foreign official flows on yields particularly challenging.

This paper revisits the issue of estimating the price impact of foreign official demand for USTs, taking new steps to explicitly address biases arising from both simultaneity and omitted factors. Existing estimates vary widely (Table 1), with prominent studies indicating that a $100 billion foreign official purchase (sale) of USTs can lower (raise) long-term yields by 13 to 68 basis points. Estimates at the upper end of this range are sizable, consistent with the 'Global Savings Glut' Hypothesis.[2] But the margin of uncertainty across different studies is wide. The extant literature traditionally controls for U.S. domestic fundamental factors and tries to overcome simultaneity by implicitly assuming that foreign official demand

[1]Foreign official institutions typically encompass state-sponsored actors such as central banks, finance ministries, and sovereign wealth funds. 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 Aizenman et al. [2021], He et al. [2021], Vissing-Jorgensen [2021].

[2]See Greenspan [2005] on the 'Interest Rate Conundrum', Bernanke [2005] 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.

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].

is inelastic. However, simultaneity bias remains if foreign official demand is even partially elastic, or if there exists other price-elastic investor segments that do respond to changing Treasury prices. And there may be additional important factors, such as Federal Reserve monetary policy and foreign yields that jointly shape U.S. yields and foreign official demand for USTs.

We show that the impact of foreign official demand for USTs on U.S. yields is substantially understated when these critical sources of endogeneity are left unaccounted for. We exploit the change in capital flow volatility following the 2008 Financial Crisis to estimate a VAR identified via heteroskedasticity of foreign official flows and U.S. yields of different maturities, controlling for both domestic and foreign factors. An identified foreign official flow shock of $100 billion moves the 5-year, 10-year, and 30-year yield by over 100 basis points in the short-run before quickly reverting to the range reported in previous studies shown in Table 1. Back-of-the-envelope calculations based on our estimated price impacts suggest that a sudden reduction of the USD share of reserves held by foreign central banks such as China or Saudi Arabia may have a substantial - if transitory - effect on U.S. interest rates.

Table 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
Warnock and Warnock (2009)	-34 to -68	12M FO flows, Treasuries+Agencies (% GDP)	1984M01-2005M05
Bertaut et al. (2012)	-13	FO holdings, Treasuries+Agencies (% debt)	1980Q1-2007Q2
Beltran et al. (2013)	-39 to -62	12M FO flows, Treasuries (% debt)	1994M01-2007M06
Beltran et al. (2013)	-46 to -50	FO flows, Treasuries (% debt)	1994M01-2007M06
Beltran et al. (2013)	-17 to -20	FO holdings, Treasuries (% debt)	1994M01-2007M06
Wolcott (2020)	-17	FO flows, Treasuries (% debt)	1985M01-2014M08
This study: OLS	-19 to -44	12-month FO flows, Treasuries (% debt)	1999M01-2018M12
This study: SVAR	-100 to -140	FO flows, Treasuries (% 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 'FO' refers to 'foreign official'.'*This study*' estimates based on benchmark-consistent official flows data of Bertaut and Tryon [2007] and Bertaut and Judson [2014]. See for a broader set of estimates: [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].

Previous estimates assume that U.S. yields are determined mainly by domestic fundamentals such as the short rate, inflation, or growth expectations of the United States. Most researchers studying the impact of foreign official demand therefore control for such domestic factors and 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 like that of other preferred-habitat investors such as pensions and insurers [Greenwood and Vissing-Jorgensen, 2018], because reserve managers do not prioritize maximizing risk-adjusted returns. By contrast, some argue that reserve managers are at least partially price-elastic [Borio et al., 2008; Chinn et al., 2022]. Even under fully inelastic foreign official demand, the fact that other investor segments do exhibit price-elastic demand for Treasuries causes the simultaneity problem to persist. Alternatively, instruments such as FX interventions or trade flows of granular countries have been proposed 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].

As the literature focuses on controlling for U.S. domestic factors, the presence of foreign factors which may jointly shape foreign official demand for USTs and U.S. yields introduces another source of endogeneity on which the literature has remained 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, the same global factors also drive foreign official UST demand because the precautionary and mercantilist motives behind foreign official UST accumulation tend to be pro-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 their currency. In addition to foreign factors, Fed monetary policy may jointly impact U.S. yields and foreign official demand for USTs, to the extent that a monetary policy shock impacts the Dollar, affecting economic conditions abroad or inducing reserve managers to intervene in FX markets. So while foreign official demand for USTs is inelastic to an extent, it also depends on the state of the global economy and on additional U.S. factors like Fed monetary policy. As a result, prevailing estimates may still be biased if they don't explicitly deal with simultaneity or the control for such factors. Our aim is to address both of these issues.

Inferring the direction of bias caused by simultaneity in realistic settings is generally difficult. But we can take an *a priori* stance based on the context of our problem under reasonable assumptions and a simple model. Assuming the causal impact of UST purchases on U.S. yields is negative, further suppose that the causal impact of U.S. yields on UST demand is positive (based on uncovered interest rate parity). In the presence of simultaneity, the estimated effect of UST purchases on U.S. yields will confound the negative true effect with the positive effect of U.S. yields on UST demand, leading the estimate to be less negative, or understated, compared to the true effect. This simutaneity bias persists in cases where foreign officials exhibit inelastic demand but there exists other price-inelastic investor segments. Simultaneity bias can therefore be economically interpreted as a case where the price impact of foreign official flows is endogenously dampened by price-elastic investors entering the market attracted by these new prevailing prices. In addition, the sign of the bias caused from omitting variables can go in either direction, and depends on the covariance between foreign official UST purchases and the omitted factor, and the covariance between U.S. yields and the omitted factor.

The baseline sample under consideration is the period January 1999 to December 2018. The period intentionally spans two decades characterized by rapid global economic and financial integration. We report both OLS and identified estimates of the impact of foreign official UST demand on U.S. yields. We start by estimating a benchmark regression following Warnock and Warnock [2009]. 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 of about 19 basis points and 44 basis points if the sample is truncated at December 2007.[3] The 19 to 44 basis point range from the regression specification is in line with prominent estimates in the literature.

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; Lewis, 2022] exploiting a well documented regime change in the pattern of global capital flows. Specifically, we exploit the change in the volatility of foreign official flows that occurred around the time of the 2008

[3]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.

global financial crisis. Our OLS estimate increasing upon truncating the sample period at the end of 2007 is consistent with an increase in foreign official flows volatility after the crisis. We further validate the structural break in volatility in three distinct ways. First, we refer to a broad literature reporting the change capital flow dynamics since the 2008 crisis. Second, we conduct variance ratio tests under a known break and find statistically significant changes in the volatility of foreign official UST flows since the crisis. Third, we test for multiple unknown structural breaks and identify a significant break in 2008. At the same time, we extend the set of traditional controls by conditioning on both domestic and foreign factors. Along with U.S. fundamental factors, we control for foreign sovereign yield factors recovered from a panel of short-term and long-term government bond yields of 19 non-U.S. advanced economies, along with Fed conventional and unconventional monetary policy. The identified estimated impulse responses show that a $100 billion foreign official UST sale or purchase causes a change in 5-year yields of 112 basis points and 10-year yields of about 125 basis points upon impact. These effects are large but transitory, falling quickly into the range reported in previous studies. We go on to estimate several VAR specifications for robustness, finding that the impact on long-term yields is substantially larger than previous estimates across all models.

Such large price impacts are not observed in practice since foreign official flows are highly endogenous and typically any effect is dampened by opposing forces. However, our results provide new insights in a quasi-controlled scenario where all else remains fixed. We see that the price impact of foreign official UST flows are consequential, and the role of foreign investors for Treasury market functioning may be more important than previous estimates imply. As a case study, we use our identified price impact estimates to assess the implications of a large, exogenous reduction in Dollar reserve assets by foreign central banks such as China or Saudi Arabia. We find the effects to be substantial. We estimate that a 10 percent reduction in Dollar reserves of China, holding roughly $3.25 trillion in FX reserves, amounts to a -$195 billion outflow, causing 5-year and 10-year yields to temporarily rise more than

200 basis points. A 10 percent reduction in Dollar reserves from Saudia Arabia, holding roughly $326 billion in FX reserves, would entail outflows of -$20 billion, resulting in 5-year and 10-year yields rising between 20 and 25 basis points. Again, we interpret the size of these effects as an extreme scenario resulting from truly exogenous, rather than endogenous, UST sales.

The rest of the paper is organized as follows. Section 2 discusses data, presents benchmark OLS estimates, and briefly discusses theoretical biases. Section 3 sets up our structural VAR identified through heteroskedasticity based on the structural break in official flows volatility since the 2008 financial crisis, and reports the main result of the paper. It also presents results from a back-of-the-envelope scenario analysis where foreign central banks exogenously reduce their Dollar share of FX reserves by 10%. Section 4 concludes. The Appendix provides extensive 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, robustness of the VAR identification assumptions, and discussion on comparing foreign official and domestic official (Fed) UST purchases.

2 Data, Benchmark OLS Estimates, and Theoretical Biases

2.1 Data and OLS specification

The baseline sample period is January 1999 to December 2018 for which we have benchmark-consistent flow data.[4] Starting in the early 2000s aligns well with modern era of globalization, the period during which the 'Global Savings Glut' also arose. We examine longer sample periods in a series of robustness checks. Following the literature, we separate foreign official

[4]The details on the data sources and transformation for all variables are in Section A1 of the Appendix.

flows from aggregate or foreign private flows. The existing literature typically estimates the following specification with OLS, based on the seminal contribution of Warnock and Warnock [2009]:

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

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

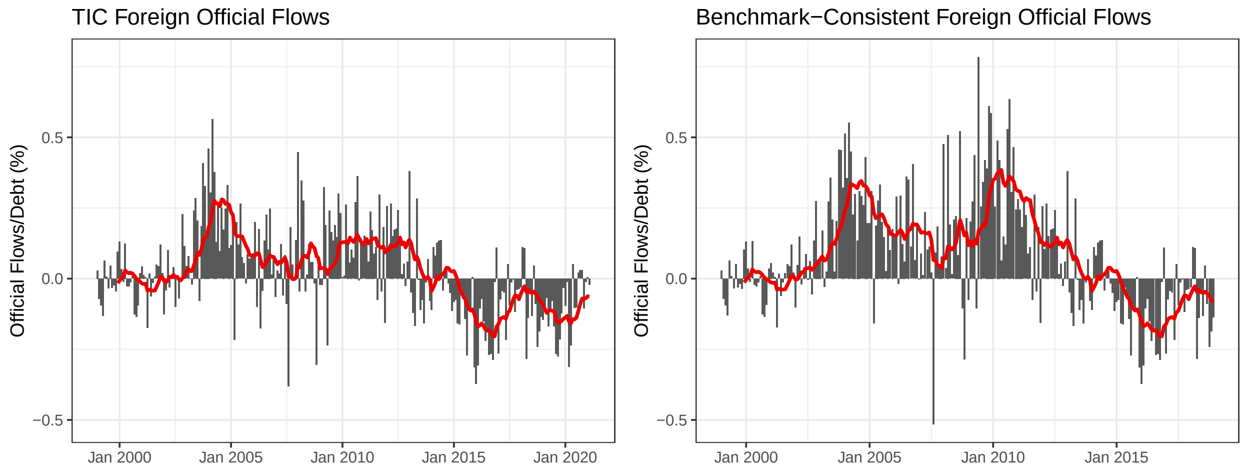
where $y_{us,t}^{10Y}$ and $y_{us,t}^{3M}$ are U.S. long and short-term Treasury yields corresponding to the 10-year and 3-month maturity, respectively.

In our estimation of this benchmark model, we exclude an autoregressive term, implying that the long-term impact does not take into account feedback effects from previous changes in the U.S. yield, although our VAR specification that follows fully allows for feedback effects. ΔFO_t denotes foreign official purchases and sales of U.S. Treasury bonds and notes, scaled by 12-month lagged U.S. marketable debt outstanding, computed as detailed in the Appendix Section A1. For interpretation purposes, we also follow the literature and 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 magnitude of the regression coefficient. Across all specifications, we use benchmark-consistent official UST flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014] because of the well-known limitations of raw TIC data, although we also consider the raw TIC data which allow us to extend the sample to February 2021 as a robustness check.[5] Benchmark-consistent and TIC flows are plotted in Figure 2, providing evidence of cyclicality and also a possible mean and variance-shift during the recovery from the global financial crisis, on which we expand below.

The specification above includes several domestic controls traditionally used in the literature: 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 (VIX_t); the structural budget surplus or

[5]TIC data cannot differentiate official flows when the transaction goes through a third-party intermediary. Therefore, results based on TIC flows can be considered a lower-bound estimate. Second, TIC data tends to overstate purchases of some securities (such as U.S. Agency bonds). Third, TIC data does not differentiate transactions from valuation effects.

Figure 2: 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. Lines are the 12-month rolling average. U.S. marketable debt outstanding is lagged 12 months. Left panel uses TIC-reported flows data and right panel uses adjusted benchmark-consistent flows following Bertaut and Tryon [2007] and Bertaut and Judson [2014].

deficit ($surplus_t$) as a percent of GDP along with an intercept term and linear time trend ($\mathbf{1}$ and t).

The specification includes 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 (1) where $L = 11$ and $\theta_l = \theta$ for all $l = 0, ...11$. Therefore (1) encompasses 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 $\sum_{l=1}^{11} \theta_l$.

2.2 Benchmark OLS estimation and omitted variable bias

Table 2 reports the estimation results.[6] Broadly speaking the results are consistent with the literature: higher short-term yields pass through to higher long-term yields, positive GDP

[6]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.

and inflation forecasts are associated with higher 10-year yields, while higher risk as captured by the VIX index and structural budget surpluses are associated with lower 10-year yields.

Table 2: 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)	0.372***	(0.033)
1Y GDP Forecast	0.017	(0.067)	0.488***	(0.100)
10Y Inflation Forecast	0.106	(0.657)	0.347	(0.608)
1Y Inflation Forecast	0.047	(0.078)	-0.057	(0.065)
VIX	-0.021***	(0.004)	0.009	(0.006)
Budget Surplus	-0.082***	(0.037)	-0.054	(0.039)
12M Foreign Official Flows	-1.017***	(0.151)	-0.348*	(0.199)
Adj. R^2	0.909		0.916	
T	266		240	
ADF Statistic	-4.798***		-5.282***	

Regression estimates from Equation 1. 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].

The coefficient reported on the foreign official flows variable (ΔFO_t) 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 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 using raw TIC flows suggest an impact from a $100 billion flow of roughly 55 basis points. These estimates align closely with several reported in the literature despite different sample periods, methods and data (Table 1). Moreover, if we re-estimate the regression using benchmark-consistent flows but truncate the sample period to end at December 2007, the estimate increases from -0.348 to -0.80, suggesting a stronger impact of

44 basis points per $100 billion flow, in line with the results of Warnock and Warnock [2009]. The weakening of the estimate following the 2008 Financial Crisis is consistent with a rise on foreign official flows volatility since 2008, a data feature we later exploit for identification. A rise in foreign official flows volatility, else fixed, weakens the estimate of θ_l by construction because the estimate of θ_l is equal to $cov(y^{10}_{us,t}, \Delta FO_{t-l})/var(\Delta FO_{t-l})$.

2.3 Theoretical Biases

While our OLS estimates are in line with the previous literature, we argue that a causal interpretation is difficult for several reasons, including model mispecification and threats of endogeneity. The regression model does not allow for dynamics between Treasury yields of different maturities or the same yields over time, hence we first extend the static OLS framework to a VAR. Within this VAR framework we aim to carefully address the endogeneity issues that remain present. In order to motivate our VAR model main specification and to help interpret our quantitative findings, it is helpful to discuss the possible sources of bias that can affect the estimates in the extant literature. We consider two main sources of endogeneity, abstracting fro measurement error. First, the simultaneity between UST yields and UST flows. Second, the omission of relevant factors that can jointly influence UST yields and foreign demand for USTs.

Simultaneity

While the object of interest is the causal impact of foreign official flows on U.S. yields, it is plausible that U.S. yields also affect foreign flows into U.S. government bond markets if investors are price-elastic. Previous studies typically address this issue by assuming that the official component of foreign flows into these markets are price-inelastic so that the demand for USTs by this class of investors is unaffected by the price of Treasuries. Tabova and Warnock [2021] provide recent evidence supporting this assumption. However, others studies, such as Borio et al. [2008] and Chinn et al. [2022], found that official reserve managers do in fact take into account prices and yields in their asset allocation decisions.

While the theoretical sign of the bias arising from this simultaneity depends on the sign of the two causal relationships, in our context, it is possible to formulate plausible priors by examining a very simple model. Assume that the causal impact of net foreign UST purchases on yields is negative **(a)** regardless of investor type, and further suppose that the reverse effect of higher yields on UST flows is positive **(b)**, which is also plausible and theoretically consistent with uncovered interest rate parity. If $cov(a) < 0$ but $cov(b) > 0$, an estimate that confounds **(a)** and **(b)** will be less negative than the true effect of **(a)**, which we wish to estimate. Under these two plausible assumptions, in our context, simultaneity should lead to an understatement of the effect of foreign official flows on U.S. yields. More formally we lay out two simple examples. First, like described above is the case where we relax the assumption that foreign official demand is inelastic, we have the classic simultaneous equation system:

$$y_t = a\Delta FO_t + e_1, \text{ and } \Delta FO_t = by_t + e_2 \tag{2}$$

where $a < 0$ and $b > 0$. It is easily shown that in the presence of simultaneity, that the bias from estimating a will be in the direction of $b\sigma_1^2/(1 - ba)$, where σ_1^2 is the variance of e_1. With $b > 0$ and $a < 0$, the estimate of a will be less negative than the true a.

Now consider the alternative case where foreign official demand is fully inelastic ($b = 0$) but there exists other price-elastic market segments:

$$y_t = a\Delta FO_t + c\Delta PR_t + e_3, \text{ and } \Delta PR_t = dy_t + e_4, \tag{3}$$

where ΔPR_t are for example, private flows. Here, all flows negatively impact the yield, $a < 0$ and $c < 0$, but private flows are elastic, $d > 0$. It can be shown that the two equation system above can be re-written as:

$$y_t = \frac{a}{1 - cd} \Delta FO_t + \frac{ce_4 + e_3}{1 - cd}, \tag{4}$$

resulting in an estimate for the price impact of ΔFO_t of $a/(1 - cd)$. With $c < 0$ and $d > 0$, the estimate will again be less negative than the true effect a despite inelastic foreign official demand, so long as there exists elastic demand from other market segments.

Omitted Variables

U.S. yields are determined by a variety of economic and financial factors which are typically included as control variables. These include, U.S. expectations of real GDP growth and inflation, proxies for financial risk appetite, and fiscal balances. On the other hand, foreign official demand for USTs is endogenously determined by underlying precautionary, mercantilist, or exchange rate smoothing motives of foreign central banks [Obstfeld et al., 2010; Jeanne and Ranciere, 2011; Dominguez et al., 2012]. These motives, however, also depend on the state of the global economy which, in turn, also can affect U.S. yields Cesa-Bianchi et al. [2020]. We therefore extend the set of controls used in the extant literature to include additional domestic factors but also foreign common factors.

The sign of the bias from omitted variables (z_t) can go in either direction depending on the covariance between the omitted variable and foreign official flows, $cov(z_t, \Delta FO_t)$, and the covariance between the omitted variable and U.S. yields, $cov(z_t, y_{us,t})$. Both foreign official flows and U.S. yields tend to be pro-cyclical with respect to global economic cycles, as we document in Section A2.3 of the Appendix. They both increase during global economic expansions and fall during contractions. Intuitively, omitted variable bias will lead to a less negative (more negative) estimate compared to the true negative effect of foreign official flows on U.S. yields when z_t covaries positively with both ΔFO_t and $y_{us,t}$. In other words, if z_t is associated with global growth which lifts both foreign official UST demand and U.S. yields, omitting z_t will lead to a less negative estimate of the impact of foreign official flows than the true effect, thereby understating the impact of foreign official flows on U.S. yields.

It's worth noting that these simple cases may not hold in practice where the data generating process is much more complex and the true model may be unknown. Moreover, it is possible that some factors introduce both omitted variable bias and simultaneity bias at the same time. For instance, Federal Reserve asset purchases can be both cyclical and introduce simultaneity if their purchases of Treasuries coincide with foreign official purchases or sales (e.g., they may be counter-parties). As a result, these theoretical exercises are meant to provide a useful framework to think about potential biases, but in practice it proves difficult to cleanly quantify and separate biases caused by omitted variables and bias caused by simultaneity, especially when they coexist.

3 Identifying a UST Flow Shock

The regression analysis in the previous section, while useful to compare estimates against benchmarks in the literature, does not control for the simultaneity between U.S. yields and foreign official purchases. To address this thorny identification issue the literature uses VARs identified via short-run zero restrictions, but existing estimates vary widely and such restrictions may not reasonably deal with the simultaneity problem.[7] In this section we address identification in novel manner by estimating a structural VAR identified through heteroskedasticity in foreign official UST purchases.

Specifically, to estimate the dynamic impact of a foreign official UST purchase or sale shock, we estimate a VAR including U.S. yields of differing maturities and (benchmark-consistent) foreign official UST flows specified as follows:

$$\mathbf{Y}_t = \boldsymbol{\beta}'\mathbf{Y}_{t-l} + \boldsymbol{\Gamma}'\mathbf{X}_t + \mathbf{u}_t, \tag{5}$$

[7] 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.

where

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

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

Here, the vector of endogenous variables, $\mathbf{Y}_t$, in addition to net foreign purchases (ΔFO_t), includes 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, that effectively de-trends these level variables while accounting for systematic Fed monetary policy.

The vector of control variables, $\mathbf{X}_t$, includes the domestic factors used in the static model: expected 1-year ahead real GDP growth, expected 1-year and 10-year inflation, the VIX index, and the structural budget surplus. As additional control variables, we include foreign yield factors recovered from a panel of short-term and long-term government bond yields of 19 non-U.S. advanced economies (described in detail in Section A2 of the Appendix). We also include the high-frequency identified Federal Reserve monetary policy shocks of Swanson [2021]. The three shocks that we include are shocks to the Fed Funds Target, Forward Guidance, and LSAP respectively. Finally, we include three dummy variables, represented by D_t, to remove major three large residual outliers that give rise to wide error bands in the bootstrap procedure used to construct the confidence set around the impulse response 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.[8]

The VAR above is estimated with 4 lags similar to Bernanke et al. [2004]. In extensive robustness analysis we find that the results are robust to both decreasing and increasing the lag length. The VAR in (5) can be viewed as a dynamic extension of the static model in (1). It is still parsimonious in its assumptions and is transparent.

[8]We also considered including dummies for the September 11, 2001 terrorist attacks and the 2013 'Taper Tantrum'. Including these additional dummies tightens the confidence bands further, but also affect the volatility of the shocks before and after the GFC.

3.1 SVAR identification through heteroskedasticity

We address our simultaneity problem by identifying the structural shock interest in the VAR above. To do so, we exploit a shift in the time series' variance of the foreign official purchases after the Lehman crash in September 2008. Identification via heteroskedasticity was initially proposed in Rigobon [2003]. Brunnermeier et al. [2021] apply it in a VAR setting. Identification solely based on short-run zero restrictions is difficult to justify within a setting like ours where asset prices (yields) and quantities (foreign official flows) often respond contemporaneously to each other. Identifying foreign official flow shocks using sign restrictions is another possibility, but there are other shocks which generate negative correlations between foreign official UST purchases and U.S. yields, for example U.S. monetary policy shocks that impact the U.S. Dollar thereby inducing foreign central banks to sell reserves to stabilize their currency. Finally, studies which rely on external instruments require that instruments identify exogeneous variation in foreign official flows. However, detailed data on commonly used instruments is often unavailable to the public (e.g. central bank interventions) or not exogenous (e.g. China's trade balance).

The basic idea of heteroskedasticity-based identification is as follows.[9] Suppose the covariance matrix of the residuals $\mathbf{u}_t$ differ before and after September 2008:

$$E(\mathbf{u}_t\mathbf{u}_t') = \begin{cases} \Sigma_1, \text{ for } t = 1, ..., t_{Sep2008} - 1 \\ \Sigma_2, \text{ for } t = t_{Sep2008}, ..., T, \end{cases} \tag{6}$$

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_{kk}, $k = 1, ..., 6$, with $K = 6$ and B is the Cholesky factor or 'impact matrix'. The relationship between covariance matrix of the reduced form residual 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

[9] See Olea et al. [2021] for a succinct treatment.

change in variances in second regime. The matrix of structural shocks, ζ_t, is identified if all elements of Λ are distinct.[10] The shocks are identified up to a column order and rotation. In our application, we are only concerned about identifying the foreign official flow shock. Therefore we only need to assume a structural break in the volatility of foreign official UST flows. We can then select the shock to the variable of interest from the set of identified shocks using procedures such as those laid out in Lewis [2021] and Lewis [2022].

3.2 Evidence of the 2008 break in foreign official flows volatility

Assuming a structural break in the variance of the UST foreign purchase time series is critical for the validity of identification through heteroskedasticity.[11] As we showed in Section 2.2, the OLS estimate of the impact of foreign official flows on U.S. yields is substantially stronger in the sample ending just before the 2008 Financial Crisis compared to the full sample that is estimated through 2018. This result is consistent with an increase in foreign official flows volatility after the crisis. All else fixed, a rise in foreign official flows volatility in the post-crisis period will increase the denominator in the computation of $\theta_l = cov(y^{10}_{us,t}, \Delta FO_{t-l})/var(\Delta FO_{t-l})$ estimate over the full sample relative to the pre-crisis sample. We provide three additional, distinct pieces of evidence supporting the presence of such a break after the 2008 Financial Crisis: (1) historical/literature-based, (2) variance tests under a known break, (3) structural break tests under unknown breaks.

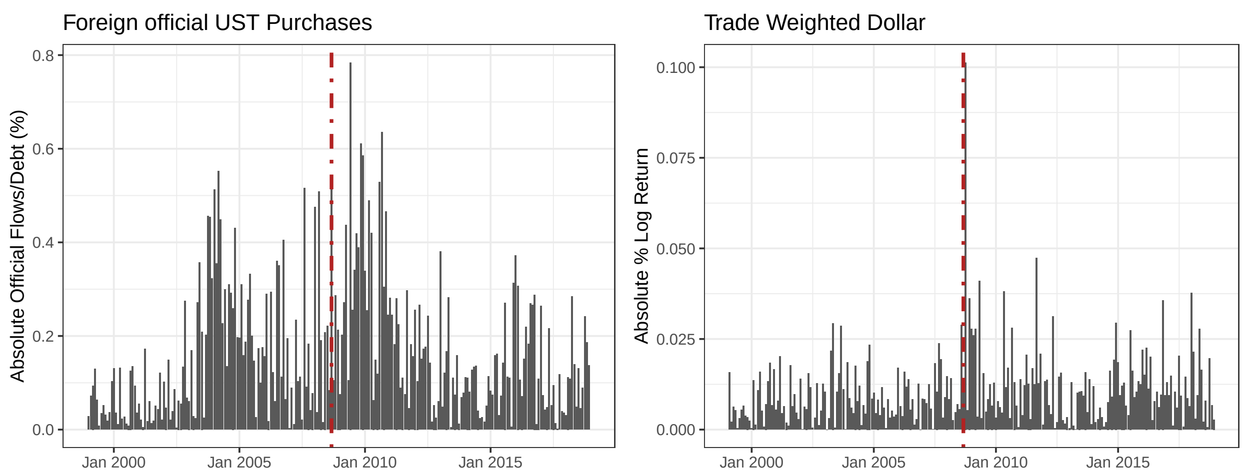
Historical context. Studies exploiting heteroskedastic-based identification typically base their credibility on external knowledge of historical events. In our case, we assume a structural break in September 2008 in coincidence with the Lehman crash because of the well-documented shift in patterns of international capital flows before and after the global financial crisis along multiple dimensions. For example, Ahmed and Zlate [2014] document

[10] Specifically, adding a second regime doubles the number of identifying equations from $(K^2 + K)/2$ to $(K^2 + K)$ while adding K new unknowns to the previous K^2 unknowns, thus the VAR is just identified.

[11] As Lewis [2021] shows, identification can also achieved under the weaker conditions that there is time-varying volatility and conditional independence of the structural shocks. As the evidence supporting our identifying assumptions is compelling, we don't need to invoke these milder conditions.

significant changes in net capital flow behavior after the 2008 crisis. López and Stracca [2021] documents that capital flows fell 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 drivers of capital flows changed from risk appetite to commodity prices, and extreme capital flow movements decreased since the crisis. Erik et al. [2020] also find that the role of the U.S. Dollar factor has gained importance since the 2008 crisis as a risk factor.

Figure 3: Absolute Foreign Official UST Flows and 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.

Variance tests (known break). Consistent with structural shifts in capital flows documented by the literature, we find that the volatility of foreign official UST flows changed since 2008 in a statistically significant manner. Figure 3 left panel plots the absolute value of foreign official UST flows, a proxy for volatility. The volatility of foreign official UST flows changed significantly in the post-2008 period. Indeed, the monthly standard deviation of foreign official flows increased about 30%, from 0.17% to 0.23%, and an F-tests for the ratio of pre/post September 2008 variances indicates that the shift is statistically significant at the 1% level (Table 3). The right panel plots monthly absolute logged U.S. traded-weighted Dollar returns. On an annualized basis, the volatility of the U.S. Dollar also increased, almost doubling, from 3.77% to 6.14% in the period after September 2008. Again, the change in variance is significant at the 1% level. These large and statistically significant regime changes in official flow volatility is robust to assuming alternative break dates (e.g., if we exclude the GFC crisis period altogether) and we find similar changes in UST flow volatility when inspecting large country UST holdings as in the case of China.

Table 3 also reports variance changes in yield variables of the VAR. Associated with the Fed Fund rate hitting the effective lower bound during the post 2008 period, the variance of U.S. yields (as spreads over the Fed Funds Rate) dropped substantially after September 2008. Post-2008 variances of yield variables are 60-90% of the pre-September 2008 period variances. With the variance of foreign official flows rising by a significant margin while the remaining variances falling assures that our variable of interest is the one that changes volatility by the largest magnitude relative to other variables, consistent with necessary condition for identification spelled out in Lewis [2021], Lütkepohl et al. [2021] and others.

Table 3: Estimated Variances of VAR Variables in Pre-Post September 2008 Regimes

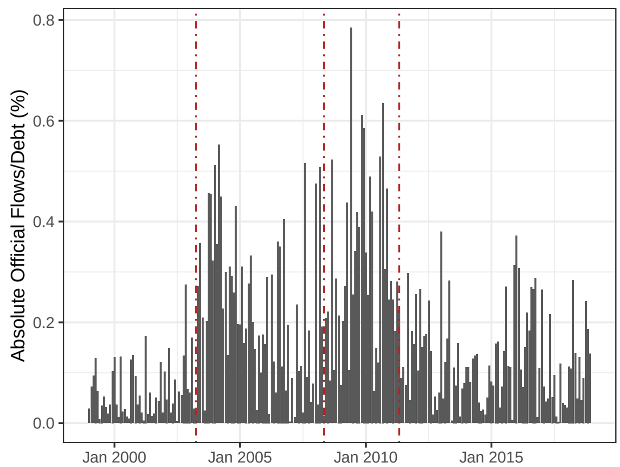
	ΔFO_t	$y_{us,t}^{3M-FF}$	$y_{us,t}^{2Y-FF}$	$y_{us,t}^{5Y-FF}$	$y_{us,t}^{10Y-FF}$	$y_{us,t}^{30Y-FF}$
Jan 1999 - Aug 2008 (R1)	0.029	0.090	0.496	1.22	1.99	3.00
Sep 2008 - Dec 2018 (R2)	0.052	0.009	0.054	0.262	0.579	0.872
F-test (R2/R1)	1.811***	0.100***	0.110***	0.216***	0.292***	0.291***

Reported sample variances within each regime period for endogenous variables in Equation 5. F-test null hypothesis is that the ratio of variances equal 1.

Structural break tests (unknown break). As a third exercise, we consider tests when the break in volatility is not known *ex ante* using the structural break testing framework of Bai and Perron [2003] which allows for the presence of multiple breaks in the series of interest. As a proxy for foreign official flows volatility, we consider the absolute value of foreign official flows as depicted in the left-panel of Figure 3 and also in Figure 4.

The structural break test detects three breaks based on the Bayesian Information Criterion (BIC): April 2003, May 2008, and May 2011. The 90% confidence bands for the may May 2008 contain the September 2008 specified break of interest. The April 2003 break date

Figure 4: Structural Breaks in Absolute Foreign Official UST Flows



Dashed vertical lines are detected breaks based on methodology of Bai and Perron [2003]: April 2003, May 2008, and May 2011.

is possibly resulting from the aftermath of several international currency crises of the late 90's and early 2000's, along with the onset of rapid foreign exchange reserves accumulation over the 2000's. While three beaks are detected, we only consider the 2008 break date for our identification, as estimation remains consistent under mispecified break dates [Sims, 2020].

3.3 Conditions for valid identification

As discussed extensively by Brunnermeier et al. [2021], a critical auxiliary assumption for credibly pursuing this identification strategy is that all VAR model parameters except the covariance matrix are stable over time; an assumption that we can investigate empirically by testing for VAR parameter stability together with a battery of other diagnostic tests on the reduced form residuals. Section A4 of the Appendix show that the reduced-form VAR residuals from the estimation of (5) are serially uncorrelated but heteroskedastic. Importantly, however, there is no evidence of VAR parameter instability, thus meeting a critical condition required for valid identification through heteroskedasticity.[12]

[12] These results are consistent with Sims and Zha [2006], who find support for VAR models with time-varying volatility over time-varying VAR coefficients.

In addition to testing for the stability of the VAR parameters, Lewis [2021] shows that credible identification may also be threatened by two other sources of identification weakness. First, the variance of the shock of interest may change by too little between regimes. Then of course there would be not enough identifying variation. This is not an empirical challenge in our context based on the evidence reported and the evidence in the literature cited above. Second, weak identification could also arises if the variances of all variables in the VAR change by the same factor [Olea et al., 2021].[13] If all variances increase by the same proportion, there is no new identifying information as the covariance matrices are just scalar multiples of each other across regimes. Moreover, it is desired that the shock of interest exhibit the largest relative change in volatility across regimes. This second condition implies that identification is not guaranteed even if there are substantial changes in variances between regimes. Referring again to Table 3, it is unlikely to be the case that all variable volatilities shift by the same factor given the increased variance in foreign official flows and the the decrease in variance across U.S. yields differentials relative to the Fed Fund rate. Following Lütkepohl et al. [2021] we formally test the VAR residuals to verify whether the volatility shifts across variables are distinct or not. Section A4 of the Appendix reports that the null hypothesis of equal variance shifts across all variable residuals is soundly rejected.

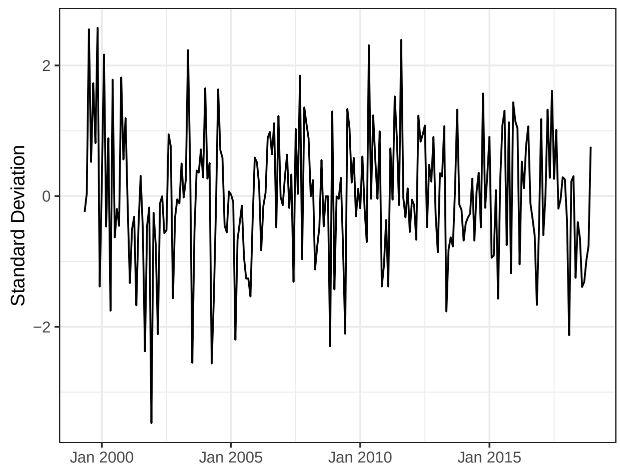
3.4 Estimated shocks and interpretation

We estimate the VAR using OLS and then recover the structural shocks following the procedure outlined in Lanne and Lütkepohl [2008], shown in Figure 5. The matrices B and Λ are initialized by maximizing the log-likelihood under the assumption of Gaussian residuals $\mathbf{u}_t$. The VAR parameters, B and Λ, are then estimated iteratively using generalized least squares. The structural shocks can be recovered using the estimated B matrix and reduced form residuals as $\zeta_t = B^{-1}\mathbf{u}_t$. As pointed out by Lewis [2021] and Brunnermeier et al. [2021], the B matrix and hence the structural shocks are identified only up to a column order and

[13]Typical examples are common shocks to the volatilities of all variables in the VAR system, such as the Great Moderation, or the global increase in volatility during the 2008 financial crisis.

rotation, so it is necessary to take an additional step in labeling and interpreting the shock of interest – in our case, the shock to foreign official flows.

Figure 5: Identified Foreign Official UST Flow Shocks



The time-series of the standardized structural foreign official flow shocks identified via heteroskedasticity from estimated VAR residuals of Equation 5.

This labeling and economic interpretation problem has no set technical solution. In practice, Lewis [2021] suggests a number of alternative strategies and steps that can be taken or have been used in previous studies. In Section A4 of the Appendix, we follow a multi-pronged strategy to pin down the shock of interest. These approaches include identifying shocks exhibiting large time-zero responses in foreign official flows (a similar strategy is followed by Brunnermeier et al. [2021]), examining whether the IRFs are theoretically consistent with the negative covariance of flows and yields, evaluating the forecast error variance decomposition (FEVD) of the different shocks (in the spirit of SVAR identification via FEVD maximization as in Volpicella [2021]), and examining the elements of the B matrix. We select the identified shock that satisfies all four criteria and this shock explains roughly 30% of the forecast error variance of foreign official flows. We are agnostic on the interpretation of the other five structural shocks in our VAR system. Note, however, that the estimated VAR is assumed to have a fundamental representation that is inevitable. One could therefore conceivably assume that the other yield differentials carry information on the level, the slope, and the curvature of the U.S. yield curve, as well as information over fundamental objectives of monetary policy such as inflation and the output gap.

3.5 Impulse Response Functions

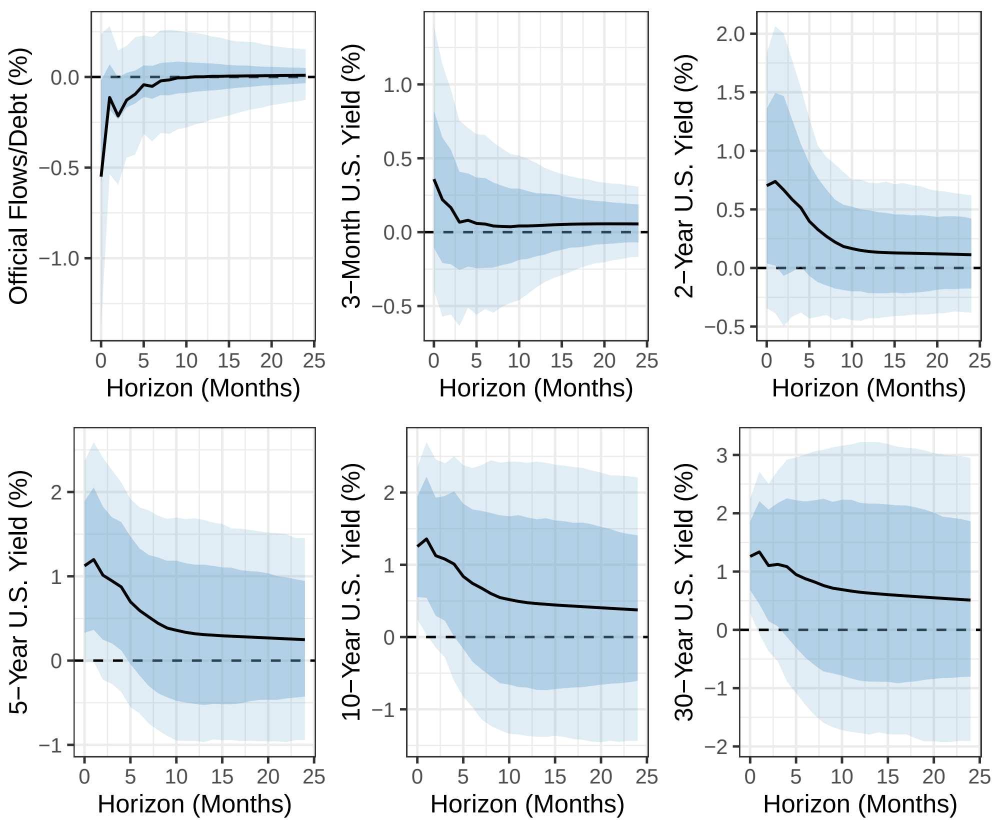
Figure 6 reports the boostrapped mean impulse response functions (IRF) to an identified official UST sale shock, together with confidence bands for foreign official flows, and the 3-month, 2-year, 5-year, 10-year, and 30-year nominal Treasury yield spreads over the Federal Funds rate. As the system is linear, IRFs to an official UST purchase are simply the same as the ones reported, but with a changed sign. Note here that, under identification through heteroskedasticity, the shape of the IRFs are identical across regimes (as long as all model parameters except the variances are stable), but the size of the effects can differ. Instead of choosing a specific regime, Figure 6 plots IRFs scaled to reflect a \$100 billion foreign official sale (or 0.55% of marketable debt basd on December 2017 debt outstanding).

Following a \$100 billion sale by foreign officials, there is no statistically significant impact on the 3-month yield. The 2-year yield rises about 70 basis points, and this effect is significant at the 68% level. The 5, 10, and 30-year yields rise by 112, 125, and 126 basis points respectively, significant at the 90% level and consistent with the existing evidence suggesting that shorter maturity yields are largely determined by domestic U.S. monetary policy and foreign purchases affect the U.S. term structure through the term premium.[14] These effects, while large, are transitory as the effects decay over several months toward zero.

Our effect estimates are substantially larger than the 13-68 basis point range reported in the prevailing literature. The magnitude of the effects are also large compared to the literature on quantitative easing (i.e. domestic official UST purchases), an issue that we discuss further in Section A5 of the Appendix. Finally, it is clearly unlikely that price impacts this large are observed in practice because foreign official flows are highly endogenous and

[14] While the larger impact on longer-term maturities is consistent with Kaminska and Zinna [2020], absense of a particularly large effect on 5-year yields is partly inconsistent with the fact that foreign official UST portfolio has an average duration of about 4-5 years [Tabova and Warnock, 2021]. Reconciling this puzzle would likely require more granular capital flow data, disaggregated by duration or at least maturity.

Figure 6: Impulse Response of U.S. Yields over the Federal Funds Rate to a Foreign Official UST Sale Shock



Impulse responses from a VAR(4) specified as in Equation 5 using benchmark-consistent foreign official flows scaled by U.S. marketable debt lagged 12 months. Light and dark shaded regions refer to 90% and 68% bootstrapped confidence intervals based on 1,000 bootstrapped samples. Solid line is the bootstrap mean IRF estimate. Shocks are scaled to reflect $100B foreign official sale of USTs.

rarely exogenous. However, this analysis helps to sheds light in a quasi-controlled setting allowing us to infer the extent of the 'unobserved' price impact that would occur absent the presence of price-elastic investors or other confounding factors.

3.6 Robustness

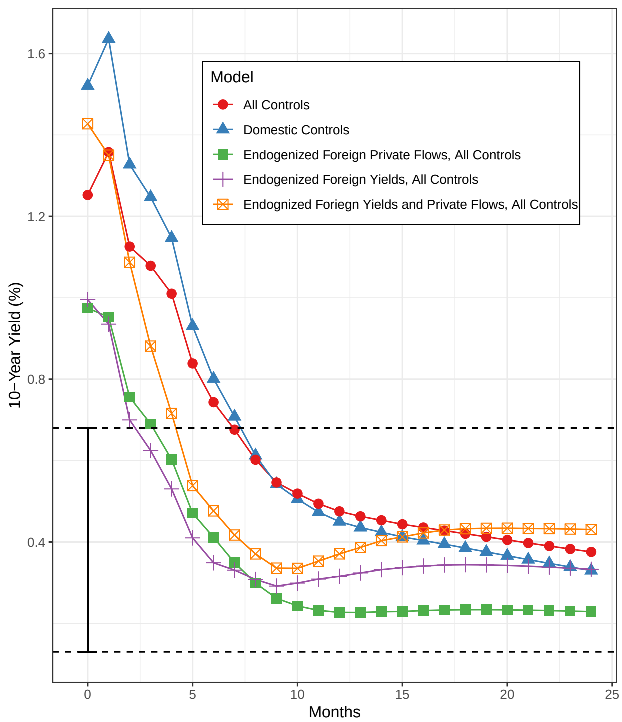
We also examine the IRFs for the 10-year yield from several additional VAR specifications: 1. SVAR with all controls (baseline in Figure 6); 2. SVAR with only domestic controls; 3. SVAR endogenizing foreign private flows; 4. SVAR endogenizing foreing yield factors instead of having them enter as controls; 5. SVAR endogenizing both foreign private flows and yield factors.

Endogenizing private flows is particularly important to account for feedback effects of private investors responding to higher yields, following a foreign official sale. Similarly, U.S. and foreign yields are likely to affect each other either via interest rate parity conditions or through the global factors that foreign yields reflect. In VAR specifications 4 and 5, the long-term and short-term foreign yield factors enter the VAR as a spread, $\mathcal{Y}_{g,t}^{10Y} - \mathcal{Y}_{g,t}^{3M}$. Figure 7 presents IRFs for the 10-year yield from the five model specifications above.

The time-zero impact across the five models range from 100 to about 150 basis points. All of these estimates exceed those from those reported in prominent studies (which range from impacts of 13 to 68 basis points). While all contemporaneous impacts are substantially large, the effects are transitory across all models, and eventually fall into the range of estimates from the literature. Figure 8 reports the full set of IRFs from model 5 which endogenizes both foreign private flows and foreign yield factors.[15] The results are qualitatively and quantitatively very similar to the baseline model (Figure 6). Further validating that a foreign official flow shock is identified, the time-zero response of foreign private flows is insignificantly different from zero, and slightly positive, implying that the identified foreign official sale of USTs is not matched by private investor purchases. The response of the foreign term spread

[15]Due to computational constraints, the confidence intervals are based on 100 bootstrapped samples.

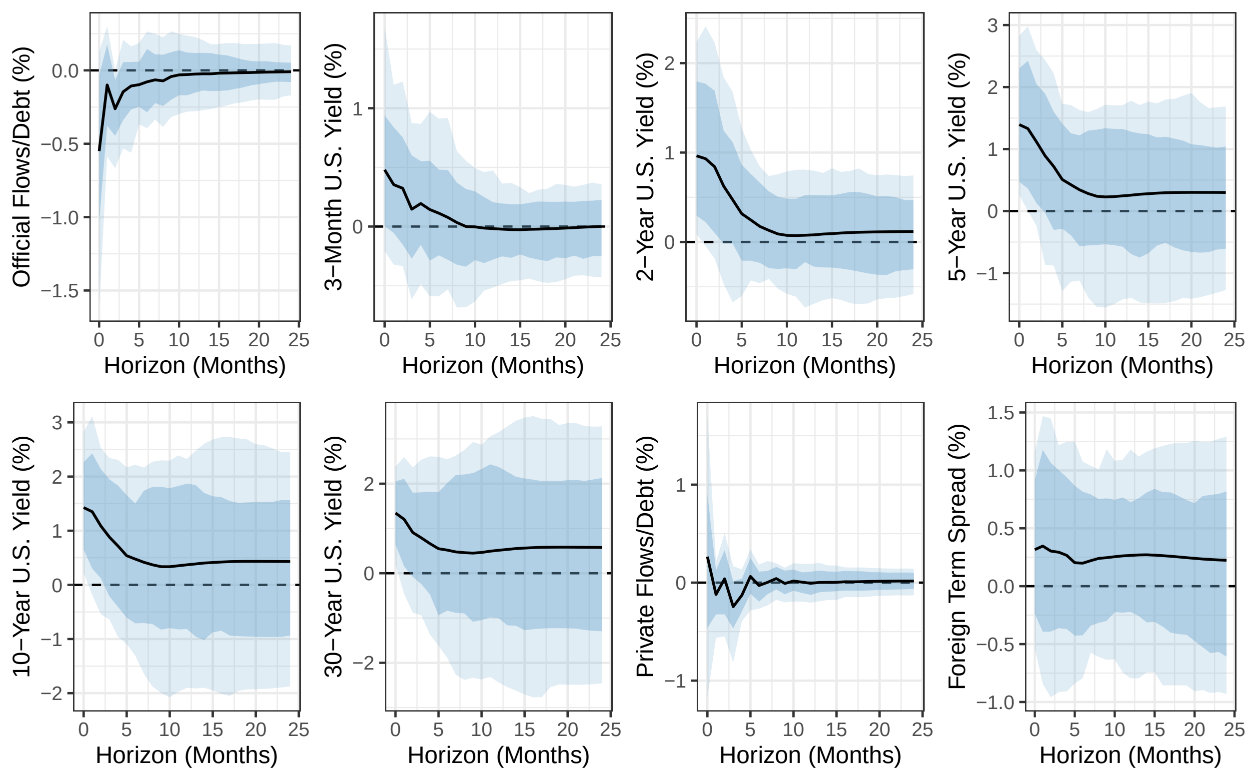
Figure 7: Impulse Response of the 10-Year U.S. Yield over the Federal Funds Rate to a Foreign Official UST Sale Shock: Alternative Model Specifications



Impulse responses from a VAR(4) specified as in Equation 5 with slight alterations outlined in Section ??. Responses are scaled to reflect $100B foreign official sale of USTs. Black point is the OLS estimated impact of -60.5 basis points from Table 2 and the error bands cover the range of estimates from previous studies [-13 bps,-68 bps].

is also insignificant, further supporting the exogenous nature of the foreign official flow shock. Over time, the foreign term spread remains positive, matching the persistent increase in the U.S. term spread which can be seen by taking the 10-year U.S. IRF and subtracting from it the 3-month U.S. IRF.

Figure 8: Impulse Response of the 10-Year U.S. Yield over the Federal Funds Rate to a Foreign Official UST Sale Shock: Model Endogenizing Foreign Private Flows and Foreign Yields



Impulse responses from a VAR(4) specified as in Equation 5 but with foreign private flows and foreign yield factors endogenized as described in Section 3.6, using benchmark-consistent foreign official flows scaled by U.S. marketable debt lagged 12 months. Light and dark shaded regions refer to 90% and 68% bootstrapped confidence intervals based on 1,00 bootstrapped samples. Solid line is the bootstrap mean IRF estimate. Shocks are scaled to reflect $100B foreign official sale of USTs.

3.7 A shift away from Dollar reserves by foreign central banks

In this section we present back-of-the-envelope calculations of the price impact on U.S. yields arising from a foreign central bank reducing its holdings of USD-denominated assets by 10 percent. Specifically we consider China and Saudi Arabia, two major holders of U.S. Dollar assets. Over the last decade, China has ranked as either the largest or second largest holder of USTs, with total FX reserves (excluding gold) reaching over \$3.25 trillion as of March 2022. Much of China's reserves accumulation is considered a byproduct of its mercantilist trade policies and exchange rate targeting monetary policy. Saudi Arabia, another country under consideration, operates a hard peg to the Dollar and also generates significant oil export revenues. Both of these features underpin the country's large holdings of Dollar reserves. Note some simplifying assumptions made in this back-of-the-envelope calculation. First, we assume that all foreign official flows are in the form of FX reserves when in actuality foreign official flows can also entail flows from other government entities such as sovereign wealth funds. Moreover, we assume that all USD-denominated assets are in the form of Treasury securities.

Table 4 reports simple calculations estimating the impact on U.S. yields if either of these countries decided to exogenously reallocate their FX reserves away from USTs by 10 percent. We assume that of total FX reserves, 60% are comprised of USTs based on the aggregate USD share of FX reserves reported by the IMF Currency Composition of Foreign Exchange Reserves (COFER).[16] An exogenous 10 percent shift out of USTs by China amounts to a -\$195 billion outflow ($\$3,250\mathrm{B} \times 0.6 \times -0.10$). Under an elasticity of 1.12 and 1.25 basis points per \$1 billion flows, 5-year and 10-year yields rise 218 and 244 basis points, respectively. The same 10 percent reallocation by Saudi Arabia results in a -\$20 billion outflow, implying the 5-year and 10-year yield rise by 22 and 25 basis points, respectively. These effects are expected to be short-lived, and can be considered extreme scenarios. In practice, large, completely exogenous sales of USTs rarely occur and it is likely the observed impact would be smaller as other market participants endogenously purchase Treasuries as yields rise.

[16] The exact USD share of global FX reserves reported by COFER as of Q4 2021 was 58.81%.

Table 4: Price Impact from a 10% Reduction of USD Reserves of China and Saudi Arabia

	China	Saudi Arabia
FX Reserves ($B, Mar 2022)	3,250	326
Assume % USD	60%	60%
10% Outflow ($B)	-195	-20
5Y yield elasticity per $1B	1.12bps	1.12bps
Contemporaneous impact on 5Y yield	**+218bps**	**+22bps**
10Y yield elasticity per $1B	1.25bps	1.25bps
Contemporaneous impact on 10Y yield	**+244bps**	**+25bps**

Assumes all USD-denominated reserve assets are U.S. Treasury securities. Assumed share of USD assets of 60% is taken from IMF Currency Composition of Official Foreign Exchange Reserves (COFER). Yield elasticities are from the VAR model estimated in Equation 5.

4 Concluding Remarks

Exploiting the change in foreign official flows volatility following the 2008 Financial Crisis, we demonstrate using a VAR identified through heteroskedasticity that the impact of foreign official demand for USTs on U.S. yields is substantially understated when critical sources of endogeneity are not accounted for. We estimate that an identified $100 billion sale of USTs by foreign officials raises the 10-year yield by more than 100 basis points in the short-run before falling within the range of 13 to 68 basis points reported in previous studies. Moreover, it is likely that effects this large are rarely observed in practice due to the highly endogenous nature of Dollar reserves accumulation by foreign official institutions.

These results bear important policy implications. The impact of UST liquidations or purchases that are truly exogenous may have a substantially larger impact on U.S. Treasury market functioning and financial conditions than previously thought. Back-of-the-envelope calculations suggest that a 10 percent shift away from Dollar reserves held by foreign central banks such as China and Saudi Arabia can substantially impact U.S. Treasury yields in the short-run.

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Online Appendix

The Appendix is organized as follows: Section A1 provides detail on the data sources and the construction of the variables used in the analysis. Sections A2 provides detail on constructing the foreign yield factors and A3 presents evidence of the common factor in foreign bond yields, and shows that foreign yield factors are not spanned by the level, slope, and curvature factors of the U.S. term structure when explaining U.S. 10-year and 3-month yields. It also provides robustness checks on factor construction and presents stylized facts connecting common foreign yield factors to global economic activity, pro-cyclical foreign official UST flows, and counter-cyclical private investor UST flows. Section A4 presents VAR diagnostics that support our approach of identification through heteroskedasticity.

A1 Data

All data used in this analysis is publicly available and collected from a variety of sources. First, to construct foreign bond yield factors, 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 gross domestic product (GDP) is also taken form the OECD for these countries. 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. Country-specific yield and GDP data are used to construct GDP-weighted foreign yield factors. 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 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 values, we use the rates paid on Japanese 3-month certificates of deposits which are available on the Federal Reserve Economic Database (FRED). The analysis also uses 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) from FRED.

Global monthly industrial production growth measures (U.S., advanced economy ex. U.S., emerging market) are taken from the Dallas Fed Database of Global Economic Indicators, which covers a core sample of 40 countries. 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 control for other determinants of U.S. bond yields beyond foreign demand, we consider variables previously 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. U.S. marketable debt outstanding is calculated as total U.S. public debt less public debt held by the Federal Reserve banks. 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). 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 (GDP growth expectations, inflation expectations, budget surpluses) are linearly interpolated to monthly frequency. 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.1: 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 Foreign Yield Factor	266	2.651	1.406	0.008	1.103	3.756	4.892
Short-Term Foreign Yield Factor	266	1.586	1.409	−0.287	0.230	2.722	4.272
1-Year 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 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 limitations. First, TIC data cannot identify an official flow 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 tends to overstate purchases of some securities (such as U.S. Agency bonds). Third, the raw data tends to contain valuation effects which need to be stripped out. For these reasons, we 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 is available up to December 2018, while the TIC raw data sample ends in February 2021. Summary statistics for the main variables are reported in Table A.1.

A2 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 is abundant.[A1] The presence of global economic forces implies the joint determination of foreign UST demand, international and U.S. bond yields. In this section, we show that 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 also 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, common foreign factors must be controlled for to avoid omitted variable bias.

As an illustration, suppose long-term U.S. yields depend linearly on both domestic and

[A1] See for example, Ilmanen [1995], Diebold et al. [2008], Obstfeld [2020]). 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], productivity and real activity [Bekaert and Ermolov, 2021], or common monetary regimes [Borio et al., 2019]. In the context of foreign demand for USTs, 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.

foreign factors along with foreign official UST flows:

$$y_{us,t}^{10Y} = \mathbf{a_1'} F_{US,t} + a_2 \Delta FO_t + \mathbf{a_3'} F_{G,t} + e_{us,t}, \tag{A.1}$$

where $y_{us,t}^{10Y}$ is the U.S. 10-year Treasury yield, $F_{US,t}$ represents a set of domestic factors such as the short-term rate and other fundamentals, ΔFO_t are UST net foreign official purchases, $F_{G,t}$ represents possibly *unobserved* foreign factors, and $e_{us,t}$ is an i.i.d. error term. 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 (A.1) under the assumption that foreign official demand is inelastic, but they exclude other foreign factors, $F_{G,t}$. So long as $F_{G,t}$ 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 any simultaneity.

A2.1 Estimating the common factors in foreign yields

Empirically, the first challenge is to reasonably approximate the foreign component of U.S. yield variation, $F_{G,t}$. 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 ($F_{G,t}$) using weighted averages of the cross-section of international long and short-term government bond yields. The idea is that bond yield co-movement captures the foreign component of U.S. yield variation. We consider only 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 U.S. Treasury substitutes (e.g. German, Japanese, and Swiss bonds).[A2] 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.

Including emerging market yields in the construction of the common foreign factor would be inappropriate for several reasons despite the fact that, Save Japan and Switzerland, most large official UST holders are emerging markets. Foreign official demand is likely associated with idiosyncratic emerging market shocks that are *uncorrelated* with global investor demand and economic conditions, a component of variation we wish to identify rather than partial out. Emerging market yields also have sizable credit risk components [Du and Schreger, 2016], and the model assumption implicit in (A.1) that $F_{G,t}$ determines U.S. yields is likely violated: U.S. and broad advanced economy conditions affect emerging markets but the reverse is less likely to drive the data. 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 synchronized with U.S. and emerging market economic conditions (Figure A.2 below).

In constructing foreign common factors from yields, it is important to note the key assumption that no underlying country bears a disproportionally large weight. This is not a plausible assumption for the United States. Rather, it is more credible that causality runs both ways with other advanced economies interest rates influencing the U.S. yields and *vice versa* [Kim and Ochoa, 2021]. 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 are less forward-looking than asset prices. 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. We deal with these issues of simultaneity by casting our empirical model in a structural VAR framework and as

[A2] 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.

a robustness check, endogenizing the foreign yield factors.

A2.2 Foreign yield factor construction

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 A3 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 while modeling forward looking variables. Alternatively, an approach which doesn't suffer from this limitation is taking the cross-section average (CSA) of yields in each period (see, 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 or economically interpretable weights to each individual country yield. Weights based on economic rationale (e.g. GDP shares) have two favorable features. First, because the weights are not estimated there is less estimation uncertainty than the PC approach for recovering unobserved factors, and second, weights can vary over time – the latter point aligning more closely with reality.[A3] Note, however, that for N large enough, the weighting scheme does not matter provided that, as we stated earlier that no individual unit dominates; a concern that we address by not including the U.S. bond yields in the cross-section average.

With CSA, weights can be economically motivated depending on the context, e.g. based on GDP shares or relative supplies of aggregate public debt. 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. Constructing both a long-term and short-term factor sufficiently accounts for the presence of a foreign 'level' and 'slope' factor in yields [Diebold et al., 2008]. 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 foreign bond yield factors as follows:

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

U.S. yields are omitted here also because they are the dependent variable of interest, and the common factors will be used as explanatory variables in the analysis.[A4] 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: $w_{i,t} = GDP_{i,t} / \sum_{i \notin us} GDP_{i,t}$.[A5] 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.

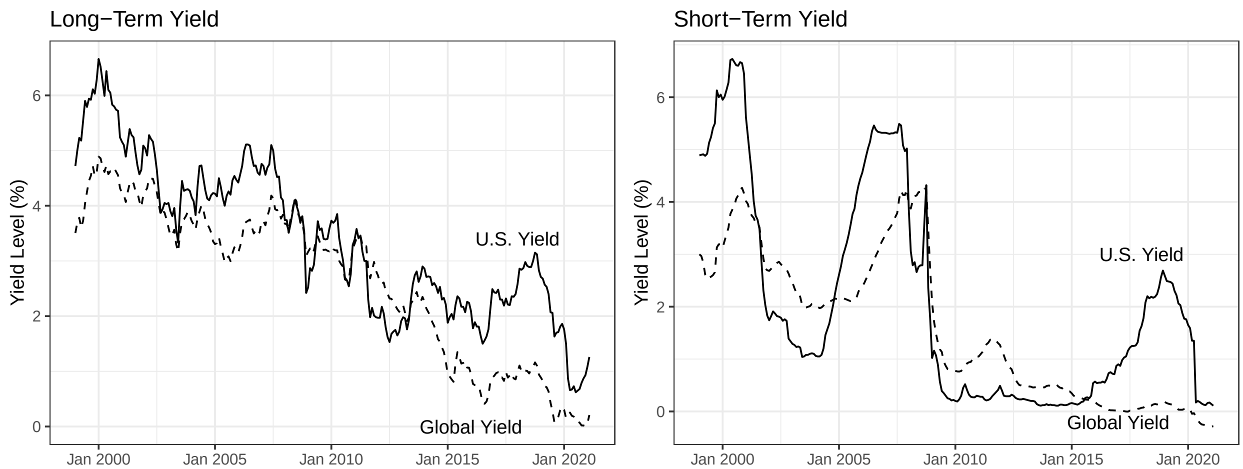
Alternatively, weights reflecting differences in the supply of national public debt stock across countries could allow for a more market-based approach. For instance, Japan's U.S. Dollar value of public debt is high even relative to GDP, and a debt-based weighting scheme

[A3] Non-estimated weights are 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.

[A4] Assuming the U.S. is a dominant (or granular unit), this implies that the world cross section of bond yields is driven by at least two common factors, the U.S. and the common foreign factor that is the focus on our analysis.

[A5] See Section A1 of the Appendix on the construction of the these weights.

Figure A.1: Common Foreign (non-U.S.) Yield Factors and U.S. Interest Rates



Solid lines are the U.S. 10-year (left panel) and 3-month (right panel) yields. Dashed lines are the common foreign 10-year (left panel) and 3-month (right panel) yield factors. Common foreign yield factors are constructed as GDP-weighted averages of non-U.S. yields as in Equation A.2.

would account for Japan's large market size. Section A3.2 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, reflecting the fact that our cross section of 19 bond yields is sufficiently large.

Figure A.1 plots GDP-weighted 10-year and 3-month common foreign yield factors along with U.S. interest rates for the same maturity. Changes in common foreign 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 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 the common component of the foreign rates.

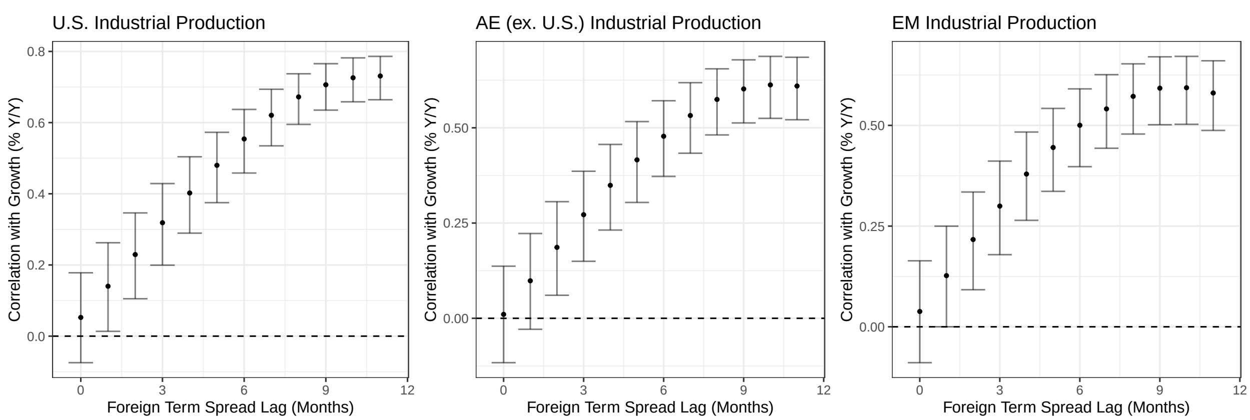
Since the start of the sample period in 1999, both short-term and long-term U.S. yields seem to diverge from these common foreign yield factors systematically before U.S. recessions: in the early 2000s amid the 'dot-com' boom, prior to the 2008 Global Financial crisis, and again prior to the 2020 pandemic 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 reverting to foreign yields and also some degree of foreign yields converging with U.S. yields.

In Section A3.1 of the Appendix, we show that these foreign common yield factors co-move with U.S. 10-year and 3-month yields *even after* controlling for traditional domestic term structure factors (level, slope, curvature), suggesting a statistically important relationship between foreign yields and U.S. yields.

A2.3 Co-movement among demand for Treasuries, U.S. interest rates, and common foreign factors

Foreign common yield factors, U.S. yields and foreign demand for USTs are intertwined. Here we want to illustrate the critical co-movements, which are a potential source of omitted variable bias as discussed above, without establishing any causal relationships among the variables of interest.

Figure A.2: Foreign Yields and their Correlation with Global Growth



The foreign yield factor term spread is constructed as GDP-weighted averages of non-U.S. yields as in Equation A.2, $\mathcal{Y}^{10Y}_{g,t-l} - \mathcal{Y}^{3M}_{g,t-l}$. Dynamic correlations and 95% confidence bands are reported, lagging the foreign term spread between $l = 0, ...11$ months against industrial production. Sample period is January 1999 to December 2018. All variables are monthly frequency. Time 0 reports the contemporaneous correlation.

For instance, optimistic global growth expectations can result in 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. Conversely, during 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, while foreign officials, mostly emerging markets, tend to liquidate international reserves to stabilize exchange rates.[A6]

To illustrate and summarize the relationships among foreign common yield factors, U.S. yields and foreign demand for USTs we focus on the difference between the 10-year and 3-month common foreign yield factors ($\mathcal{Y}_{g,t}^{10Y} - \mathcal{Y}_{g,t}^{3M}$), which we call the foreign yield factor term spread.[A7] Figure A.2 shows that this foreign yield factor term spread is correlated with global economic conditions over different horizons. This "common" foreign term spread is significantly and positively associated with current and future industrial production growth rates globally: in the U.S., in advanced economies (ex. U.S.), and in emerging market economies.

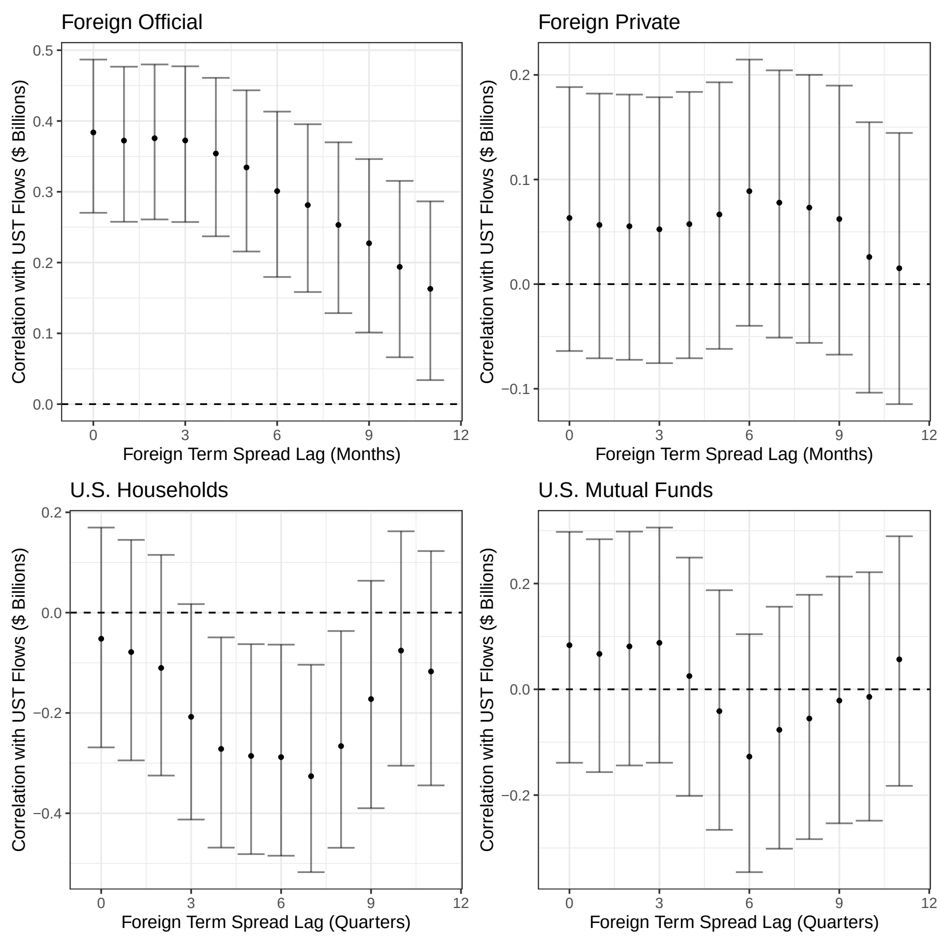
At the same time, Figure A.3 shows that the common foreign term spread is also significantly and positively associated with current and future demand for USTs by foreign *officials* (pro-cyclical UST demand) but substantially less so with foreign *private* UST demand. In contrast, the common foreign term spread is also negatively associated with current and future demand for USTs by domestic U.S. households (including hedge funds), consistent with these investor classes being a source of counter-cyclical UST demand. Like foreign private demand, U.S. mutual fund demand co-moves more weakly with the foreign common term spread. Finally, Figure A.4 illustrates the co-movement between the foreign common term spread, foreign official USTs purchases, and the U.S. term spread. The correlation coefficients are positive, sizable, and statistically significant.

Taken together, the co-movements documented are consistent with the hypothesis that

[A6] Dominguez et al. [2012] and Ahmed [2020] document evidence of cyclical reserves management driven by global shocks, while He et al. [2021] and Vissing-Jorgensen [2021] report large sales of USTs during crises.

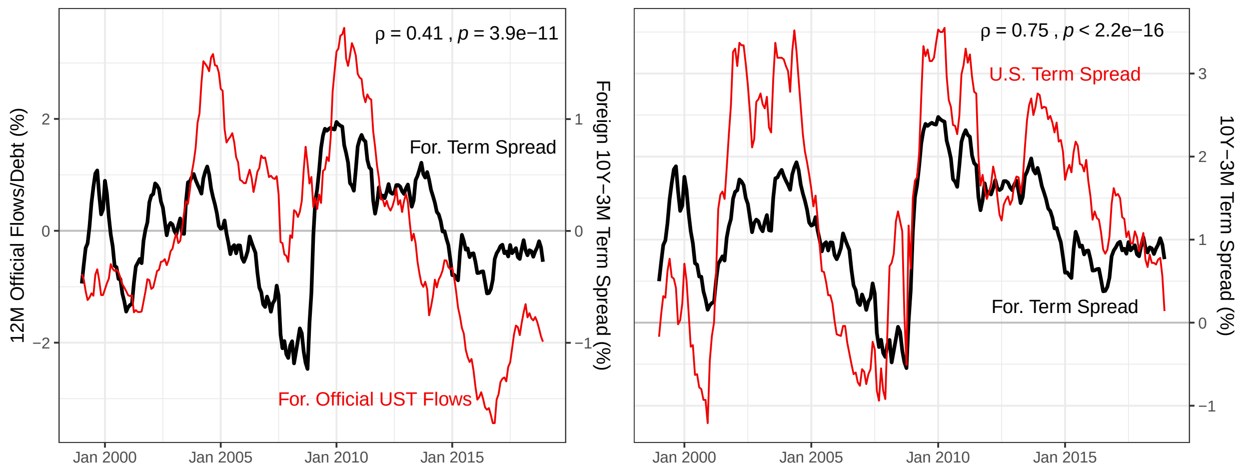
[A7] Looking at common factors in short and long-term yields separately depicts a similar more complex picture.

Figure A.3: Foreign Yields and their Correlation with Foreign and Domestic UST Demand



The foreign yield common term spread is constructed the difference between the GDP-weighted averages of non-U.S. yields in Equation A.2, i.e., $\mathcal{Y}^{10Y}_{g,t-l} - \mathcal{Y}^{3M}_{g,t-l}$. Dynamic correlations and 95% confidence bands are calculated by lagging the foreign term spread $l = 0, ...11$ months against measures of UST purchases. Foreign official and private flow data are 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 are quarterly. Time 0 is the contemporaneous correlation.

Figure A.4: Foreign common Term Spread, Foreign Official UST Purchases (left), and U.S. Term Spread (right)



The thick black line is the foreign common term spread ($\mathcal{Y}_{g,t}^{10Y} - \mathcal{Y}_{g,t}^{3M}$). Left panel overlays the foreign term spread with foreign official UST flows (both series are de-meaned). Foreign official flow data are adjusted benchmark-consistent flows (Bertaut and Tryon [2007] and Bertaut and Judson [2014]) and calculated as a 12-month rolling sum. Right panel overlays the foreign term spread with the 10-year less 3-month U.S. term spread, and series are not de-meaned. Sample period is January 1999 to December 2018.

foreign officials reserve managers accumulate USTs pro-cyclically as rates rise internationally, while private investors behave counter-cyclically buying USTs when global rates are falling and foreign officials are selling. Omitting foreign common factors, which positively co-vary with foreign official UST demand and U.S. yields, can therefore bias estimates of the effect of foreign official UST purchases on U.S. yields from negative toward zero or positive estimates.

A3 Common Foreign Yield Factors, Factor Strength, and Stylized Facts

We now present formal evidence that our 3-month and 10-year foreign yield factors capture the cross-sectional dependence governing international interest rates. Principal Component Analysis (PCA) reveals strong evidence of commonality among these two cross sections of yields, with Table A.2 providing detailed statistics.

Consider the long-term (10Y) bond yields. The first principal component (PC) of the bond yield levels accounts for 91% of the variation across the 19 non-US bond markets, and the first PC of bond yield changes captures 66% of the cross-country variation. The second PC explains 6% and 9% in levels and changes, respectively. While the variation explained by the second PC is not trivial, the large proportion of variance captured by the first PC motivates considering at least one common factor for both long-term and short-term bond yields. The results are very similar for short-term yields. Consistent with the PC analysis, the average of the 171 pair-wise correlations between the 19 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 40, respectively for long-term yield level and changes). Again the results are about the same for the short term-yields Formal (CD) tests for cross-sectional dependence [Pesaran, 2021] in bond yield levels and changes all strongly reject the null hypothesis of no cross-sectional dependence for both long and short-term yields.

Table A.2: Evidence of Factor Structure in Foreign Bond Yields

	10Y Yields		3M Yields	
	Levels	Changes	Levels	Changes
1st PC % Variance Explained	0.91	0.66	0.90	0.71
2nd PC % Variance Explained	0.06	0.09	0.05	0.07
Average Pairwise Correlation	0.90	0.62	0.88	0.66
CD Test Statistic	192.77***	131.49***	187.87***	140.06***

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 19 non-U.S. 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.

We also can formally test factor strength, as proposed by Bailey et al. [2020], by determining the proportion of country long-term or short-term yields in which a foreign factor significantly explains country-specific yield variation, while adjusting for multiple hypothesis testing. The factor regression used to assess factor strength in foreign long-term and short-term bond yields are:

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

and

$$y_{i,t}^{3M} = a_i^{3M} + B_{1i}^{3M} y_{i,t-1}^{3M} + \pi_i^{3M} \bar{y}_{-i,t}^{3M} + e_{i,t}^{3M}, \tag{A.4}$$

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 an estimate of the foreign common yield factor constructed as the equal-weighted cross-section average of 10-year yields from the other 18 non-U.S. advanced economies *excluding* country i (denoted $\bar{y}_{-i,t}^{10Y}$). The regression for short-term yields is similar with just lagged short-term yields and the 3-month foreign cross-section average ($\bar{y}_{-i,t}^{3M}$) on the right-hand side.

The coefficients of interest are the respective factor loading, π_i^{10Y} and π_i^{3M}, in the respective regressions. Factor strength can be measured by the proportion of π_i estimates which are statistically significant, out of the total 19 estimates. Critical values which indicate significance are adjusted to account for multiple testing using the Bonferroni correction.[A8] For long-term yields, at the 10% and 5% levels (after adjusting for multiple testing), all but two estimates of π_i^{10Y} across the 19 country regressions are statistically significant (89.5%). At the 1% level, 16 of the 19 of factor loading estimates (84.2%) are significant. For 3-month yields, at the 10% and 5% levels (after adjusting for multiple testing), 13 estimates of π_i^{3M} across the 19 country regressions are statistically significant (68.4%). At the 1% level, 12 of the 19 (63.2%) of factor loading estimates are significant.

According to Bailey et al. [2020], 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 these results, both the common factors in foreign long-term and short-term bond yields are confirmed as semi-strong. In unreported regressions, we also show that factors remain strong when we do not control for lagged 10-year yields and 3-month yields in the 10-year yield regressions or lagged 3-month yields in the 3-month yield regressions. We conservatively conclude from

[A8] The Bonferroni adjustment can be considered a standard case of the more general multiple testing correction described in Chudik et al. [2018].

this preliminary evidence that the common factors in foreign long-term and short-term bond yields are not weak, but semi-strong. This conclusion remain valid if we include U.S. yields in the analysis, increasing the set of yields to 20 from 19, thereby testing for the presence of a *global* factor in international long- and short-term bond yields.

A3.1 Foreign yields and U.S. yield curve factors

As it is well known, fluctuations in the U.S. yield curve are explained well by small number of factors, usually three: the 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 common foreign yield factors help explain U.S. yields beyond the traditional domestic yield curve factors. In other words, we ask whether the foreign common yield factors in this paper are spanned by the domestic U.S. 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 our findings confirm that result. To test whether the common foreign 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 auxiliary regressions for the 10Y and 3M U.S. yield:

$$y^{10Y}_{us,t} = \beta^{10Y}_0 + \beta^{10Y}_1 level_{us,t} + \beta^{10Y}_2 slope_{us,t} + \beta^{10Y}_3 curve_{us,t} + \beta^{10Y}_4 \mathcal{Y}^{10Y}_{g,t} + e_{us,t} \tag{A.5}$$

$$y^{3M}_{us,t} = \beta^{3M}_0 + \beta^{3M}_1 level_{us,t} + \beta^{3M}_2 slope_{us,t} + \beta^{3M}_3 curve_{us,t} + \beta^{3M}_4 \mathcal{Y}^{3M}_{g,t} + e_{us,t}, \tag{A.6}$$

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 common foreign yield factor for long and short-term interest rates are denoted $\mathcal{Y}^{10Y}_{g,t}$ and $\mathcal{Y}^{3M}_{g,t}$, respectively. A significant coefficient on β_4 would suggest that the foreign yield factors are not spanned by traditional U.S. yield curve factors, and that the international yields are important to consider when modeling variations in U.S. yields. To avoid collinearity, in (A.5), the 10-year U.S. yield is excluded when extracting the factors from the yield curve, and similarly in (A.6) we exclude the 3-month U.S. yield the term structure factors because these are the specific maturities that we are trying to explain. Validating our estimation procedure, note that the estimated factors omitting the 10-year yield are nearly perfectly correlated with the estimated factors obtained omitting the 3-month yield, and there is little to no practical difference between the two variants.

Table A.3: 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)
U.S. Level	1.284***	(0.018)	1.356***	(0.090)
U.S. Slope	0.414***	(0.015)	0.840***	(0.023)
U.S. Curvature	0.038***	(0.008)	0.125***	(0.033)
10Y Foreign Yield	0.035**	(0.015)		
3M Foreign Yield			0.381***	(0.073)
Adj. R^2	0.997		0.983	
T	266		266	

Regression estimates from Equations A.5 and A.6. 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 have zero mean and unit variance while the common foreign factors are in 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.5 the 10-year U.S. yield is excluded when extracting the factors from the yield curve. Similarly, for Equation A.6, the 3-month U.S. yield is excluded when extracting the factors. Sample period: January 1999 to February 2021.

Table A.3 reports ordinary least squares (OLS) estimates for (A.5) and (A.6). For both the U.S. 10Y and 3M yield, corresponding foreign yield factors $\mathcal{Y}_{g,t}^{10Y}$ and $\mathcal{Y}_{g,t}^{3M}$ enter the regressions significantly after controlling for domestic U.S. level, slope and curvature factors, despite the fact that the yield curve factors alone can explain over 99% and 96% of the variation in 10Y and 3M U.S. yields, respectively. This confirms that, indeed, common foreign yield factors contain information relevant for U.S. interest rates that is not spanned by traditional U.S. yield curve factors. Of course, this evidence cannot speak to the direction of causality, and assumes that in equilibrium individual U.S. yields are a function of both domestic yield curve and common foreign factors. U.S. yields likely spill over to foreign yields given the sizable role of the United States economy and its financial sector and, at the same time, common factors in foreign yields influence U.S. yields. Additionally, both foreign common factors, $\mathcal{Y}_{g,t}^{10Y}$ and $\mathcal{Y}_{g,t}^{3M}$, can be included in the regressions (A.5) and (A.6). When this is done, both long and short-term foreign 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 foreign factors in explaining U.S. yields.

A3.2 Alternative common factor weights

This section compares different approaches to recover common foreign bond yield factors from the cross-sections of advanced economy long and short government bond yields. The benchmark approach uses GDP weights, which vary over time and takes into account economic size differences across countries. For robustness, we also consider public debt weights that 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. An equally-weighted average is period-by-period across the 19 non-U.S. country yields is also considered, called cross-section averaging (CSA). Finally, we also estimate a PCA-based factor which assigns country weights based on estimated PCA factor loadings, but keep those weights constant over time.

Tables A.4 and A.5 report correlation coefficients among alternative measure of the common factors in foreign yields for the long-term and short-term yields, respectively. The correlations are very high, most above 0.99, implying that the foreign yield factors are robust to alternative weight schemes.

Table A.4: Common Foreign Factor in 10-Year Yields: Alternative Weighting Schemes

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

Correlations between foreign common long-term bond yield factors constructed with alternative weighting schemes as in Equation A.2. GDP refers to the GDP-weighted average. Debt refers to the 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.5: Common Foreign Factor in 3-Month Yields: Alternative Weighting Schemes

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

Correlations between foreign common short-term bond yield factors constructed with alternative weighting schemes as in Equation A.2. GDP refers to the GDP-weighted average. Debt refers to the 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.

A4 VAR Residual Diagnostics and Identification

This section presents statistics on the reduced form VAR residuals from the estimation of VAR in (5).

Table A.6: Higher Moment Summary Statistics of Reduced-Form VAR Residuals

	ΔFO_t	$y_{us,t}^{3M-FF}$	$y_{us,t}^{2Y-FF}$	$y_{us,t}^{5Y-FF}$	$y_{us,t}^{10Y-FF}$	$y_{us,t}^{30Y-FF}$
1st-Order Autocorrelation	0.01	0.15	0.03	0.05	0.06	0.04
SW Normality Test (p-value)	0.01	0.00	0.00	0.04	0.12	0.07
Kurtosis	4.00	12.45	4.27	3.58	3.71	3.76
Skewness	0.03	1.43	0.41	0.27	0.20	0.10

Summary statistics on reduced form VAR residual series from Equation 5. SW Normality test refers to the Shapiro-Wilk test where the null hypothesis is that the variable is normally distributed.

Table A.6 shows that the VAR residuals exhibit little to no meaningful autocorrelation. To examine the empirical distribution of reduced form VAR residuals, we consider several statistics. First we conduct Shapiro-Wilk tests for normality on individual residual series. For all series, the null of normally distributed residuals is rejected. We also estimate higher moments, kurtosis and skewness for each individual residual series. Kurtosis for all variables is larger than 3 suggesting the presence of fat tails, while except for the 3-month yield series, the other residual series do not exhibit substantial skewness. In addition to individual tests, multivariate tests for normality, skewness, and kurtosis reject the null hypothesis of normality. We also provide histograms of the residuals in Figure A.5. A critical assumption for our identification strategy through heteroskasticity is the presence of time-varying volatility. We do find that multivariate ARCH-LM tests strongly reject the null hypothesis of no heteroskedasticity in the residuals (Chi-square statistic of 2637.7 on 2205 degrees of freedom). These results are consistent the rejection of the variance ratio tests in Section 3.

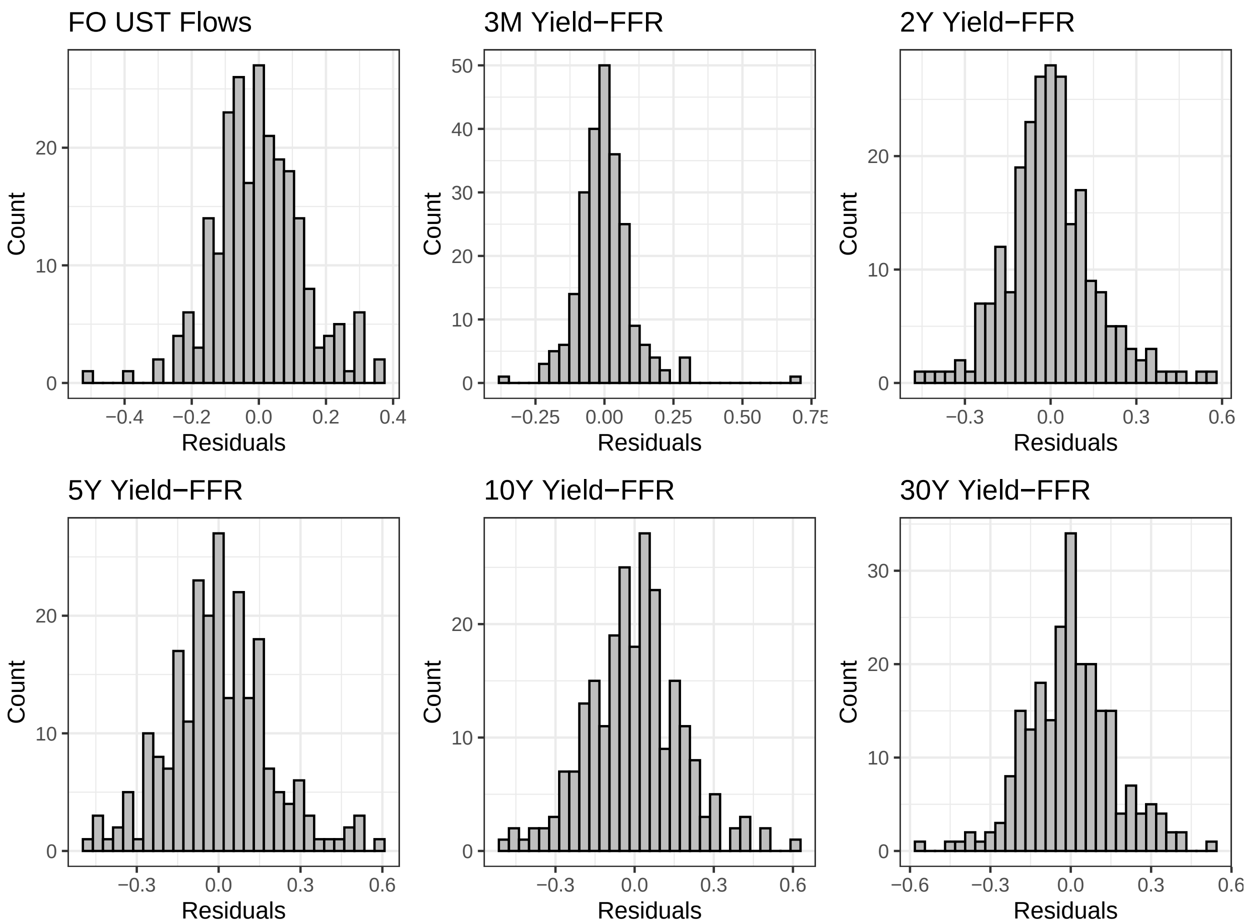
Another crucial assumption is that the VAR coefficients themselves remain stable despite the shift in volatility. Whether or not parameters of the VAR are stable in the presence of time-varying volatility can be tested by examining parameter stability of each VAR equation, by testing the path of cumulative standardized residuals [Ploberger and Krämer, 1992]. Figure A.6 resporst CUSUM plots and shows that there is no evidence of instability or breaks in VAR parameters since no series extends beyond the region within the 95% confidence bands.

Finally, identification is also threatened by the risk of all VAR variable variances change by the same scaling factor between regimes. We use the joint Wald test following Lütkepohl et al. [2021],

$$H_0 : \lambda_1 = \lambda_2 = ... = \lambda_6, \tag{A.7}$$

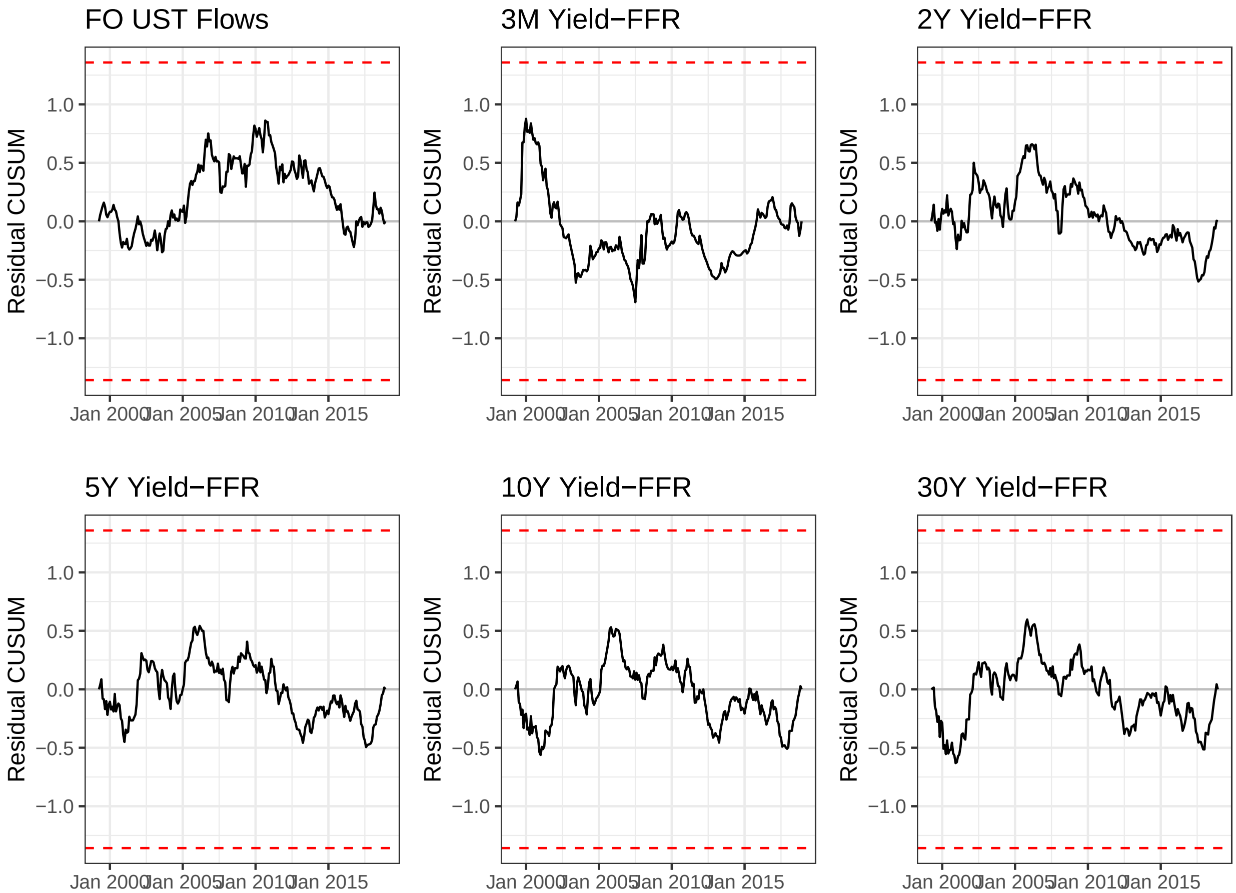
to test whether the shift in volatility (using the eigenvalues of Λ) of all variables is indistinguishable from one another. A Wald statistic of 56.64 on 20 degrees of freedom soundly rejects the null of equal variance shifts. That is, our results support the case where there is at least one distinct regime shift in volatility across our VAR variables.

Figure A.5: Histograms of Reduced-Form VAR Residuals



Histograms based on reduced form VAR residual series from Equation 5.

Figure A.6: CUSUM Test for VAR Parameter Stability



Solid lines indicate cumulative standardized VAR residual series from Equation 5. Red dashed lines are the 95% confidence boundaries. Evidence of instability or breaks in the parameters occurs if the cumulative residual series moves outside the bands.

To summarize, the reported evidence shows that the VAR residuals are uncorrelated, but perhaps not normal and they appear heteroskedastic. Moreover, we find no evidence of parameter instability or a common equal shift in variances between regimes. These results, therefore, are supportive of the statistical conditions needed for valid identification via heteroskedasticity.

A4.1 Labeling the identified foreign official flow shock

To label our shock of interest, we follow several of the approaches suggested by by Lewis [2021] and Brunnermeier et al. [2021]. The estimated B matrix is:

$$\hat{B} = \begin{bmatrix} -0.079 & 0.022 & -0.005 & 0.083 & -0.023 & -0.080 \\ 0.043 & 0.110 & 0.073 & 0.017 & 0.044 & 0.069 \\ -0.017 & 0.008 & 0.235 & 0.004 & -0.036 & 0.066 \\ -0.055 & 0.027 & 0.231 & -0.028 & -0.077 & 0.113 \\ -0.093 & 0.032 & 0.199 & -0.050 & -0.051 & 0.128 \\ -0.109 & 0.012 & 0.131 & -0.039 & -0.023 & 0.163 \end{bmatrix}. \tag{A.8}$$

We first examine which IRFs exhibit large time-zero responses to foreign official flows, our shock of interest. We also assess which IRFs are consistent with the theory-implied causal responses of the yields to a foreign official flow shock. We then examine the FEVDs in the spirit of Volpicella [2021] FEVD-based identification to examine which shock explains the most variance in foreign official flows. We also check elements of the B matrix.

Our identified foreign official flow shock satisfies the following conditions: (i) It produces large, time-zero responses in foreign official flows; (ii) it produces theory-consistent IRFs where yields increase when foreign officials sell USTs; (iii) the shock explains roughly 30% of the forecast error variance of foreign official flows; (iv) it is one of just two candidate shocks with consistent B matrix elements implying the structural negative relationship between flows and yields. Specifically our shock represents column 6 of B, which has a large negative loading in the first element (corresponding to foreign official flows) but positive elements on yields. The shock corresponding to column 6 is the only identified structural shock that satisfies all criteria (i) through (iv).

A5 Comparing with 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; Rebucci et al., 2022], although some studies find substantially larger effects. These estimates are substantially smaller than the 100+ basis point effect on the 10-year yield we report from the VAR estimates here in Section 3.

There are several reasons why domestic and foreign official purchases of USTs might affect U.S. bond yields differentially. The first issue is the difference in the research design of these empirical analyses. Effects of QE are typically examined within an event-study framework while the effects of foreign official purchases are examined within a regression or VAR framework using a continuous measure of foreign official purchases. The event-study nature of analyzing QE makes it difficult to disentangle the quantity effect of QE from the announcement or information effects. Moreover the high-frequency nature of QE studies allows for a precise estimate of the effect over a very short window. By contrast, the low-frequency nature of foreign official flows data allows for relatively less precise estimates, but these estimates capture effects over longer horizons.

Another important issue lies in the endogeneity of QE programs. The Federal Reserve tends to announce bond purchases when bond market liquidity is low or evaporating. 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', finding that endogeneity can severely bias the estimated price impact of foreign official UST purchases.

The economic implications of foreign and domestic official asset purchases may be starkly different. While domestic asset purchases are swaps of asset of different duration, foreign flows entering the U.S. economy provide incremental resources and purchasing power [Kohn, 2016]. Beltran et al. [2013] stresses other 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 ultimately 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 may be associated with inflation uncertainty, which would put upward pressure on yields and offset some of the downward pressure brought on by Federal Reserve asset purchases.