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

Rashad Ahmed† Alessandro Rebucci‡

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**Abstract**

This paper shows that the price impact of foreign official (FO) purchases or sales of U.S. Treasuries (USTs) is about twice as large as previously reported in the literature once critical sources of endogeneity are addressed. We also show that prevailing estimates of this price impact suffer from omitted variable bias when foreign government bond yields and Federal Reserve policies are not controlled for. By exploiting changes in the volatility of FO flows and U.S. yields after the 2008 Global Financial Crisis, we identify a FO flow shock via heteroskedasticity in a structural VAR. We estimate that a $100B flow shock moves the 5-year, 10-year, and 30-year yields by more than 100 basis points on impact, compared to the 19-44 basis points range that we estimate by assuming FO flows are price inelastic and without controlling for foreign yields and Fed actions. Our findings suggest that FO sales of USTs played a critical role during the March 2020 episode of Treasury market turmoil, and that even a small reduction in the Dollar's share of China's reserves could have a significant impact on U.S. long-term interest rates.

**Keywords:** Foreign official flows; Global savings glut; Identification via heteroskedasticity; Interest rates conundrum; International reserves; U.S. Dollar; VARs.
**JEL Classifications:** E43, E44, F21, F30, G10

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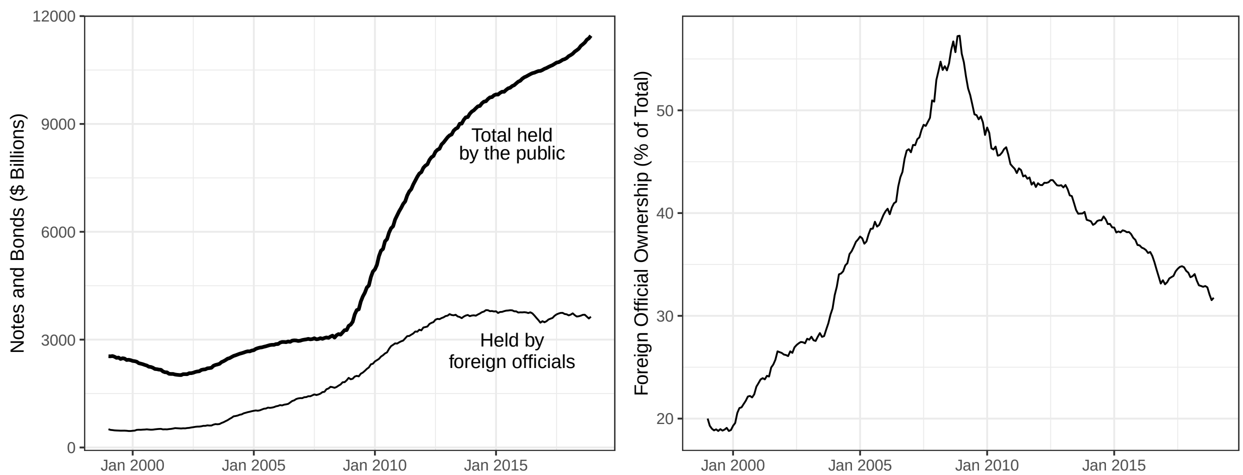
†U.S. Department of the Treasury, Office of the Comptroller of the Currency. E-mail address: rashad.ahmed@occ.treas.gov.

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

# 1 Introduction

Foreign official holdings of U.S. Treasuries (USTs) by government entities such as central banks, finance ministries, and sovereign wealth funds represents a significant share of the total U.S. public debt outstanding (Figure 1). In U.S. Dollar value (USD), foreign official (FO) holdings of USTs have increased from less than half a trillion in the early 2000s (18-20% of the total held by the public) to $3.5-4 trillion in the early 2010s (more than 55% of the total). Thereafter, FO holdings have fluctuated around this high level in USD, declining as a share of the total with fiscal policy and monetary policy increasing primary issuance and Federal Reserve holdings of Treasuries.

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




Left panel: Total U.S. Treasury notes and bonds held by the public excluding T-bills, along with the holdings of the same securities by foreign official institutions. Right panel: Foreign official holdings of notes and bonds excluding T-bills as a percentage of the total held by the public. Foreign official holdings based on the benchmark-consistent data of Bertaut and Tryon [2007] and Bertaut and Judson [2014].

A large share of FO holdings in total debt outstanding has important implications for monetary, and financial policies, as the academic and policy debates on the so-called "Interest Rate Conundrum", the "Savings Glut Hypothesis", and the Treasury market liquidity illustrate.[1] The 2008 Global Financial Crisis (GFC), the Treasury market freeze in March

[1]See, for example, Greenspan [2005] on the 'Interest Rate Conundrum', [Bernanke, 2005] on the 'Global Savings Glut' Hypothesis, Vissing-Jorgensen [2021] on the Treasury market distress in March 2020.

2020, and the consequent introduction of the Dollar swap lines with foreign central banks also reflect the financial stability risks associated with concentrated ownership of USTs by foreign officials or other classes of investors.[2] As a result, a large literature, starting with the contributions of Bernanke et al. [2004] and Warnock and Warnock [2009], attempts to estimate the impact of foreign official flows on U.S. yields.

In this paper, we revisit the estimation of the impact of FO flows on U.S. yields by identifying a FO purchase or sale shock via a heteroskedasticity-identified vector autoregression (VAR) while controlling for two previously overlooked factors: foreign government bond yields and Federal Reserve policies. Identification exploits a shift in FO flows and Treasury yield volatilities across different maturities after 2008. We find that the contemporaneous impact of a FO purchase or sale is much larger than previously estimated, reporting that a \$100B flow shock moves the 5-year, 10-year, and 30-year yield by more than 100 basis points on impact, compared to the 19-44 basis points range that we estimate in OLS regressions by assuming FO flows are price inelastic and without controlling for foreign yields and Fed actions. We also find that the estimated effects are economically significant. For example, our estimates imply that a 1% shift away from USD assets by China would lead to a 24.4 basis point increase in U.S. yields, and that FO sales of USTs may have accounted for roughly 60 basis points of the March 2020 jump in Treasury yields – abstracting from the effects of the Fed and other market participants.

The critical challenge that one faces trying to estimate the price impact of a quantity shock is endogeneity. In our setting there are at least two possible sources of endogeneity. First, demand for Treasuries affects yields but yields can also shape the quantity demanded, the more so the more price-elastic the investor segment. Second, as we document in Appendix A2, FO flows are pro-cyclical and associated with international economic forces that also affect Treasury yields, introducing confounding factors in the relationship between foreign demand and U.S. bond yields. These sources of endogeneity – simultaneity and omitted

[2] Official U.S. estimates suggest total foreign UST sales during the COVID-19 shock in March 2020 of about \$300 billion, approximately half of which are attributed to foreign government institutions.

Table 1: SELECTED ESTIMATES OF A FO PURCHASE CHANGE ON U.S. TREASURY YIELDS (IN BASIS POINTS)

| Study | Impact | Measurement | Sample Period |
|---|---|---|---|
| Bernanke et al. [2004] | -66 | Japanese Foreign Exchange 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 |

The table reports selected estimates of the impact of a FO purchase change on UST yields. Estimates can differ across studies due to differences in the scaling variable, the sample period, the modeling approach, time horizon of the estimated multipliers (typically contemporaneous or one-year), the outcome variable (10-year yield, 5-year yield, real or nominal yield, term premia, mortgage rates, etc.), or the specific Treasury securities considered. The the scaling variable (GDP, debt, etc.) used to recover these impacts is the average over the sample period used in each study. Because the scaling variables trend upward, the estimated impact of a $100B purchase is likely to be smaller than reported in more recent years. The data frequency is monthly unless otherwise noted.

variables – make consistent estimation of the price impact of FO flows on U.S. yields particularly difficult. Not surprisingly, existing estimates of the price impact of FO flows vary widely, 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 (Table 1).[3]

Inferring the direction of bias caused by simultaneity in realistic settings is generally not possible. Nonetheless, it is useful to take an *a priori* stance based on reasonable assumptions and a simple model. Assume for instance that the causal impact of UST purchases on U.S. yields is negative, while the causal impact of U.S. yields on UST demand is positive. In this case, the estimated effect of FO flows on U.S. yields will confound the negative true effect with the positive impact of U.S. yields on UST demand, leading to less negative estimates, thereby understating the true price impact of FO flows. As we show in the paper, this simultaneity bias persists in cases where FO demand is inelastic but there exists other price-elastic investor segments in the market.

Endogeneity could also arise from omitted variables introducing confounding factors in our context. The sign of these biases can go in either direction, depending on the covariance

[3]Other, less comparable, studies include: Kitchen and Chinn [2011], Kaminska and Zinna [2020], Zhang and Martínez-García [2020].

between the omitted factor and FO flows, and that between the omitted factor and U.S. yields. Consider for instance the omission of global economic conditions. Global booms are accompanied by rising safe rates, alongside growing export demand and capital inflows. Foreign central banks respond by accumulating international reserves to stem appreciating exchange rates or by building precautionary buffers. Therefore, global economic conditions are positively correlated with both U.S. yields and FO demand for USTs. Omitting such a factor would cause the estimated effect of FO flows on U.S. yields to be understated.

In our empirical analysis, we start by estimating a standard regression specification by OLS, following Warnock and Warnock [2009]. This model is estimated with the benchmark-consistent FO flows data of Bertaut and Tryon [2007] and Bertaut and Judson [2014] over the period January 1999 to December 2018. Estimating this regression under the assumption that FO flows are price inelastic, as typically done in the extant literature, we find that a $100 billion foreign official sale of USTs is associated with a rise in U.S. 10-year yields of 9 basis points on impact and 19 basis points over one-year, increasing to 44 basis points when we end the sample period in 2007.[4] This 19-44 basis point range is in line with the estimates in the existing literature shown in Table 1 that vary between 13 and 68 basis points.

We then extend this baseline regression model to an SVAR identified through heteroskedasticity following Rigobon [2003], Brunnermeier et al. [2021], and Lewis [2022]. Identification exploits a shift in the volatility of FO flows and U.S. yields occurring around the time of the 2008 GFC. U.S. yield volatility significantly fell by different degrees across the maturity spectrum after 2008, as conventional monetary policy hit the zero lower bound constraint and unconventional monetary policy started to affect the term premium. At the same time, foreign official flow volatility increased after 2008.

We document and validate this structural break in volatility by using test for known and unknown break points. Specifically, we first conduct variance ratio tests under a known

[4] For ease of interpretation, we refer to the effects of FO flows on yields in terms of either purchases or sales throughout the paper, but the effects are symmetric. In fact, we find no evidence of asymmetric effects or FO purchases versus sales.

break point, finding statistically significant changes in the volatility of FO flows and U.S. yields since the GFC in 2008. We then test for multiple unknown breaks and identify a significant break point in 2008. As we note in the paper, these shifts align well with the changes in the patterns of global capital flows, the U.S. Dollar factor, and interest rates after 2008 highlighted by Ahmed and Zlate [2014], López and Stracca [2021], Forbes and Warnock [2021], and Du et al. [2022] among others. Finally we conduct a battery of tests that reject the null hypothesis that the documented change in volatilities are proportional across variables.

In our empirical analysis, we also extend the set of traditional controls to include foreign government bond yields and Federal Reserve policies. We control for foreign yields by entering them as common factors. We control for Fed policies with the the monetary policy shocks series of Swanson [2021]. We find that controlling for either variable in our literature-consistent OLS regression model increases significantly the estimated impact of a FO flows on U.S. yields, in line with our priors on the sign of the omitted variable bias. We incorporate these variables into our SVAR framework as well.

The identified estimated impulse responses from our SVAR show that a $100 billion FO sale or purchase causes 112 and 125 basis point changes in the 5-year and the 10-year yields on impact, respectively. These impacts are much larger than those reported in the extant literature but fall within the range of the estimates reported in previous studies within five to six months. These estimated impacts are also significantly larger than the impact of Fed purchases typically found in the QE literature, possibly because foreign flows entering the U.S. economy provide incremental resources and purchasing power rather than swapping assets with differing duration.

Next, for robustness, we go on to estimate several alternative VAR specifications. We consider VARs that control only for domestic variables (as in the extant literature) and specifications in which foreign *private* flows and/or foreign yields enter as endogenous variables. We find that our results are remarkably robust, with the contemporaneous impact of a FO flow shock at least 30 % larger than previous estimates across all considered specifications.

To assess the economic significance of our findings, we estimate the possible price impact of FO sales of USTs during the Treasury market freeze in March 2020, and also consider the implications of a reduction in the share of Dollar reserves in total reserve assets by China or Saudi Arabia. We estimate that a 1 percent reduction in the share of Dollar reserves of China amounts to a -$19.5 billion outflow from the Treasury market under the assumptions made, causing the 5-year and 10-year yields to rise more than 20 basis points. A 1 percent reduction in the share of Dollar reserves from Saudia Arabia, in contrast, would entail an outflows of only -$2 billion, resulting in 5-year and 10-year yields rising between 2 and 2.5 basis points. Our estimates also suggest that FO sales may have caused U.S. yields to rise by roughly 60 basis points during the acute phase of the March 2020 Treasury market stress episode, abstracting from the offsetting impact of private and Federal Reserve purchases which also took place during that time.

**Related Literature** As we noted earlier, a large existing literature focused on the impact of FO flows on U.S. yields. To address the simultaneity problem between FO flows and U.S. yields, most previous studies rely on the critical identifying assumption that FO demand is price inelastic. Some studies handle this by simply arguing that FO demand, unlike the demand from other sectors, is inelastic like that of other preferred-habitat investors such as pensions and insurers [Greenwood and Vissing-Jorgensen, 2018]. While some motives behind FO accumulation and management of USTs may be less price-sensitive (precautionary and mercantilist), others like foreign exchange rate intervention are clearly cyclical and driven by interest rate differentials. Direct inspection of this hypothesis also suggests that reserve managers are at least partially price-elastic [Borio et al., 2008; Chinn et al., 2022; Arslanalp et al., 2022]. In this paper, we show that even under fully inelastic FO demand, the fact that flows of other investor segments, such as foreign private flows, are price-elastic can give rise to endogeneity. Yet other studies propose to use FX interventions or trade flows as instruments to isolate variation in FO demand [Bernanke et al., 2004; Beltran et al., 2013].

But again the exclusion restriction is violated if the instruments are cyclical – as we show that they are in this paper – a limitation also acknowledged by Bernanke et al. [2004]. In this paper we propose an identification strategy that does not rely on inelastic FO demand and it is robust the cyclical properties of such demand.

Identification of the price impact of FO flows is also challenging because of omitted variable concerns. Previous estimates assume that U.S. yields are determined mainly by domestic U.S. fundamentals such as the short rate, inflation and growth expectations. In this paper, we show that foreign government bond yields co-move with both FO flows and U.S. yields, likely introducing an additional source of endogeneity. The paper also shows that Fed monetary policies may also jointly impact U.S. yields and FO flows, to the extent that unconventional monetary policy shocks are counter-cyclical or impact the Dollar, affecting economic conditions abroad and inducing reserve managers to intervene in FX markets. So, while FO demand for USTs may be relatively price inelastic, it still depends on the state of the global economy and on Fed policies – factors also affecting U.S. yields. As a result, prevailing estimates are likely to be biased if they don't explicitly control for such factors. As far as we are aware, this is the first paper that considers how foreign yields and U.S. monetary policy shape the interaction between FO flows and U.S. yields.

To deal with both the simultaneity and omitted variable problem, we consider a VAR framework identified via heteroskedasticity, a strategy first pursued in Rigobon [2003] and extended to a VAR setting in Brunnermeier et al. [2021] and Lewis [2022]. Identifying FO flow shocks by imposing short-run restrictions on the VAR is inadequate in our setting because of the simultaneous nature of flows and yields. Sign-restrictions also prove less than ideal since the negative causal relationship of foreign official flows on yields can be generated by other shocks, and particularly monetary policy shocks. Instrumental variable techniques are constrained by the limited availability of data on intervention and official reserve management. While as far as we know this is the first to consider identification via heteroskedasticity of a FO flow shock, this approach has been used before in macro-finance, particularly for the estimation of the effects of monetary policy shocks on the stock market Rigobon and Sack [2003], the effects of monetary policy on interest rates and inflation [Nakamura and Steinsson, 2018], and the macroeconomic consequences of financial shocks [Brunnermeier et al., 2021].

The paper also relates to the nascent literature on the Treasury market turmoil in March 2020 (Aizenman et al. [2021], He et al. [2021], and Vissing-Jorgensen [2021]). It provides price-based evidence complementing the quantity-based analysis in Vissing-Jorgensen [2021] that FO sales played a significant role in that episode of Treasury market distress. 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. [2021a] argue that these properties are particularly appealing to foreign official investors, which can drive U.S. interest rates away from fundamentally-justified levels. Consistent with this view, we provide causal evidence of a large impact of FO flows on long-term U.S. interest rates that can account for the wedge between U.S. yields and the Fed Funds rate as in Greenspan's 'Conundrum'. Finally, the paper naturally relates to the large literature on the price impact of the Federal Reserve QE policy. We show that the price impact of FO flow shocks is significantly larger than typical QE estimates and discuss several reasons why this might be the case.

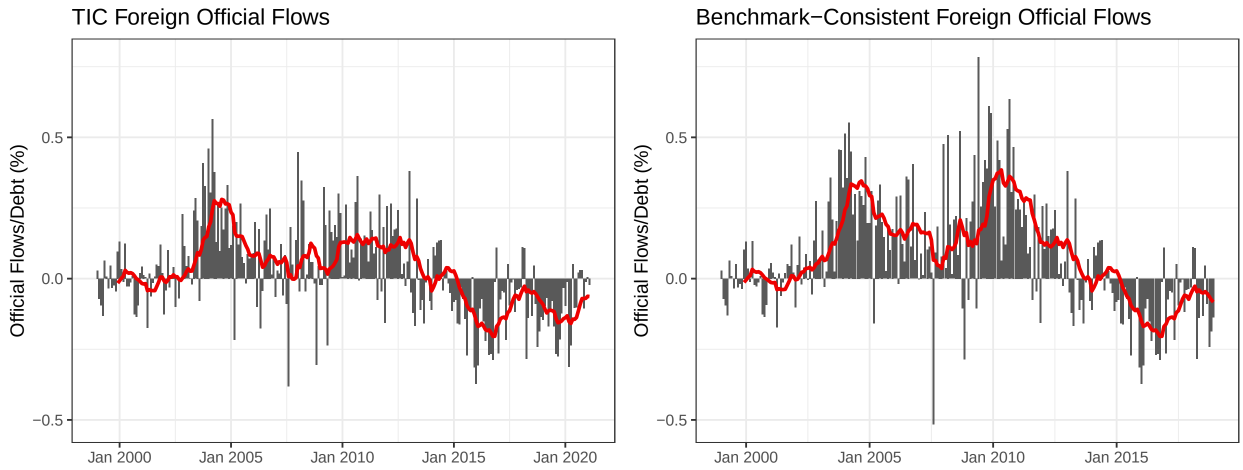
The rest of the paper is organized as follows. Section 2 discusses the data. Section 3 presents benchmark OLS estimates and discusses theoretical biases which may be present. Section 4 sets up our structural VAR identified via heteroskedasticity. Section 5 reports and discusses the main estimation results of the paper. Section 6 reports robustness analyses. Section 7 concludes. The Appendix provides details on (i) the sources and definitions of the data, (ii) the construction and properties of the foreign yield factors, and (iii) additional diagnostics on the VAR residuals supporting the identification assumptions made.

## 2 Data

The data used in the empirical analysis is publicly available from the sources described in Section A1 of the Appendix. Following the literature, we split total (or aggregate) foreign flows into official (FO) and private flows (FP), focusing our study on FO flows. Raw 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 has some well-known limitations. First, TIC data cannot identify an official flow when the transaction goes through a third-party intermediary. Thus, TIC flows represent a lower-bound estimate. Second, TIC data tends to overstate purchases of some securities such as U.S. Agency bonds. Third, the raw data also include valuation effects which need to be adjusted. For these reasons, in all our main specifications, we use *benchmark-consistent* (BC) flows that combine more accurate data from the Treasury's SHLA annual survey with the TIC SLT monthly report as proposed in Bertaut and Tryon [2007] and Bertaut and Judson [2014]. The data sources and transformations for all other variables used in the analysis are in Section A1 of the Appendix, together with a battery of summary statistics reported in Table A.1.

The BC flow data is available up to December 2018, while the TIC data sample ends in February 2021. So, our baseline sample period is January 1999 to December 2018. Starting the analysis in the early 2000s aligns well with the modern era of globalization, the period during which the 'Global Savings Glut' arose. Ending it in 2018 has the advantage to exclude the COVID-19 period. In a series of robustness checks, we also consider shorter and longer sample periods. BC and TIC flows are plotted in Figure 2, providing evidence of cyclicality and a possible mean and variance-shift during the recovery from the GFC, on which we expand below.

Figure 2: Monthly FO Net Purchases of U.S. Treasury Notes and Bonds as a Share of Marketable Debt




The bars are monthly purchases/sales. The Lines are 12-month rolling averages. Marketable debt outstanding is lagged 12 months. The left panel plots TIC flows data. The right panel plots adjusted benchmark-consistent flows (BC) data.

# 3 Benchmark OLS Estimates and Theoretical Biases

In this section, we first specify a benchmark model consistent with the approach typically adopted in the extant literature. The model is estimated by OLS under the assumption that FO flows are price inelastic and hence exogenous to U.S. yields using BC flows. Then, we discuss possible sources of endgeneity due to simultaneity and omitted variables.

## 3.1 Modeling the Impact of FO Flows on U.S. Yields

Based on the influential contribution of Warnock and Warnock [2009], the literature typically estimates some variation of the following specification by OLS, under the assumption that FO flows are price inelastic:

$$y^{10Y}_{us,t} = \phi_1 y^{3M}_{us,t} + \sum_{l=0}^{L} \theta_l FO_{t-l} + \boldsymbol{\beta'}\mathbf{X}_t + \epsilon_{us,t}, \qquad (1)$$

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

where $y_{us,t}^{10Y}$ and $y_{us,t}^{3M}$ are the U.S. 10-year and the 3-month maturity Treasury yields, respectively, $FO_t$ refers to net FO purchases and sales of U.S. Treasury bonds and notes (i.e. the first-difference of the stock of USTs held by foreign official institutions), and $\mathbf{X}_t$ is a vector of control variables.

We estimate this model with BC flow data. In our implementation of this model, we scale FO flows by 12-month lagged U.S. marketable debt outstanding, computed as detailed in the Appendix Section A1, rather than GDP as in Warnock and Warnock [2009]. Scaling FO flows by GDP does not change the results. Our benchmark specification includes $L$ lags of the FO purchases. The literature usually considers a 12-month rolling sum of FO flows, with the coefficient interpreted as the 'long-run impact' of a FO flow shock. The model (1) is more general, allowing for different coefficients on each lag of $FO_t$, and encompasses both the 12-month rolling sum typically estimated in the literature (a special case where $L = 11$ and $\theta_l = \theta$ for all $l = 0, ...11$) as well as the contemporaneous impact obtained from the coefficient $\theta_0$. Our specification above also excludes the lagged-dependent variable, implying that the dynamic short-term impacts coincide with the long-term ones. Thus, the estimated one-year impact of a FO flow shock on the 10-year yield is given by $\sum_{l=1}^{11} \theta_l$. But again our inferences and conclusions are unchanged when we include this term in the model and adjust the computations of the dynamic impact multipliers accordingly.

In the baseline specification, the vector of controls $\mathbf{X}_t$ includes the domestic controls typically used in the literature: the expected 1-year real GDP growth ($\Delta GDP_t^{E[t+1]}$); the 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 Federal budget surplus or deficit ($surplus_t$) as a share of GDP, along with the intercept term and a linear time trend ($\mathbf{1}$ and $t$).

In two alternative specifications, which we estimate and report to illustrate the scope for omitted variable bias in this standard setting, $\mathbf{X}_t$ includes variables that control for for foreign government bond yields and U.S. monetary policy shocks, as we do in our VAR analysis in Section 4 of the paper. We control for foreign yields by including in $\mathbf{X}_t$ short-term (3-month) and long-term (10-year) common foreign yield factors, constructed as GDP-weighted averages of non-U.S. sovereign yields from 19 advanced economies as described in more detail in Section A2 of the Appendix. We control for U.S. conventional and unconventional monetary policies using the three series of innovations in Swanson [2021], corresponding to policy rate, forward guidance, and large scale asset purchase (LSAP) innovations.

Table 2: Benchmark OLS Estimates

| | *Dependent Variable*: 10Y U.S. Yield | | | |
|---|---|---|---|---|
| | Benchmark-Consistent Flows | | Including Omitted Variables | |
| 3M U.S. Yield | 0.372*** | (0.033) | | |
| 1Y GDP Forecast | 0.488*** | (0.100) | | |
| 10Y Inflation Forecast | 0.347 | (0.608) | | |
| 1Y Inflation Forecast | -0.057 | (0.065) | | |
| VIX | 0.009 | (0.006) | | |
| Federal Budget Surplus | -0.054 | (0.039) | | |
| $FO$ | -0.348* | (0.199) | | |
| $FO$ $(l = 0)$ | -0.156 | (0.166) | | |
| $FO$ (Controlling for Foreign Yields) | | | -1.108*** | (0.145) |
| $FO$ (Controlling for Fed Shocks) | | | -0.983*** | (0.169) |
| Adj. $R^2$ | 0.916 | | | |
| $T$ | 240 | | | |
| ADF Statistic | -5.282*** | | | |

The table reports OLS estimates of Equation 1. Standard errors are adjusted for heteroscedasticity and autocorrelation. *, **, *** correspond to 10%, 5%, and 1% significance level, respectively. The first $FO$ coefficient and its standard error refer to the sum of the coefficients on $FO_{t-l}$ with $l = 0, ...11$, while $FO$ $(l = 0)$ refers to the contemporaneous coefficient on $FO_l$ with $l = 0$. The other two $FO$ coefficients reported are the same as FO but estimated from a specification as Equation 1 in which we control for either Foreign Yields or the U.S. monetary policy shocks as we discuss in details in Section 4. The The Augmented Dickey-Fuller (ADF) test on the regression residuals rejects the null hypothesis that the residuals are nonstationary in all three specifications. The sample period is January 1999 to December 2018 and the data frequency is monthly.

Table 2 reports the OLS estimation results for this baseline specification.[5] Broadly speaking, the results are consistent with those reported in the extant literature summarized in Table 1: short-term yields pass through to long-term yields; positive GDP and 10-year inflation forecasts are associated with higher 10-year yields, although the co-movement with inflation expectations is not significant statistically; while higher federal government structural fiscal surpluses are associated with lower 10-year yields, even though this association is not precisely estimated.

[5]All estimated specifications are tested for stationarity of the residuals via Augmented Dickey-Fuller (ADF) tests, which strongly reject the null of unit root in all cases.

The first $FO$ coefficient that we report (-0.348) refers to the 12-month sum of estimated $\theta_l$ coefficients ($l = 0, ..., 11$), which we call one-year impact. The sign of this multiplier is negative but not strongly significant statistically.[6] This estimate aligns well with several other reported in the literature despite different sample periods, estimation methods and data as summarized in Table 1. The estimate implies that a one-percentage point FO net sale leads to a cumulative 34.8 basis point increase in the 10-year yield within 12 months. Assuming marketable U.S. debt of $18 trillion as in 2017, this means that a $100B sale of USTs by foreign official holders would lead to a 19 basis point rise in the 10-year yield after 12 months.

In one of several robustness checks that we conduct, we re-estimate the baseline specification above ending the sample period in December 2007, before the peak of the GFC. Over this shorter sample period, the estimated one-year impact changes from -0.348 to -0.80 (with a standard error of 0.276, significant at the 1% level), implying that a 44 basis point change per $100B purchase or sale, in line with the results of Warnock and Warnock [2009]. This is evidence of a weakening of the impact of FO flows on U.S. yields effect after the 2008 crisis, possibly consistent with a rise in foreign official flows volatility, a data feature that we will exploit for identification purposes in the VAR analysis that follows. In fact, higher FO flows volatility, all else equal, weakens the estimate of $\theta_l$ by construction as $\theta_l$ is equal to $cov(y^{10}_{us,t}, FO_{t-l})/var(FO_{t-l})$. Of course, this result could also reflect a decline in the estimated covariance, a possibility that we discuss and rule out in Section 5.

Table 2 also reports a second coefficient on the contemporaneous value of the FO flow variable, denoted $FO$ ($l = 0$). We report this estimate as a term of comparison with our estimated VAR effects in Section 5. This coefficient is about a third of the one-year impact and is not statistically different from zero. Next, we report two more $FO$ coefficients on the right hand side of the table, corresponding to two alternative specifications in which $\mathbf{X}_t$ includes variables that control for foreign government bond yields and U.S. monetary

[6]Note that standard errors are adjusted for both heteroscedasticity and autocorrelation and hence they are conservative.

policy shocks. In both instances, we find much larger and much more precisely estimated coefficients than in the baseline specification that, as we show below, are consistent with the presence of omitted variable bias. Note here that, of the three Fed policy shocks that we control for, the LSAP innovation accounts for the large majority of the $FO$ coefficient change from -0.348 to -0.983.

## 3.2 Potential Sources of Endogeneity

While our OLS estimates above are in line with the previous literature, in this paper, we argue that a causal interpretation of this evidence is threatened by endogeneity concerns arising from potential simultaneity and omitted variable issues. Before illustrating how we propose to deal with these problems, in the rest of this section, we discuss these two possible sources of endogeneity and the theoretical sign of the associated biases that will help us interpret our quantitative findings. We focus first on the simultaneity between U.S. yields and FO flows. We then discuss the omission of foreign yield factors and U.S. monetary policy shocks that can jointly influence U.S. yields and FO flows.

### 3.2.1 Simultaneity

Our parameter of interest is the causal impact of FO flows on U.S. yields, but U.S. yields could also attract FO flows into U.S. government bond markets if investors are price-elastic. Many studies assume that the official component of total foreign flows into the U.S. government bond market is price-inelastic, with Tabova and Warnock [2021], Jiang et al. [2021b], and Fang et al. [2022] providing recent evidence supporting this assumption. However, others studies, such as Borio et al. [2008], Chinn et al. [2022], and Arslanalp et al. [2022] argue that official reserve managers do in fact take into account prices and yields in their asset allocation decisions.

The theoretical sign of the bias arising from such simultaneity depends on the sign of the two causal relationships. In our context, it is plausible to assume that $(a)$ the causal impact of net FO purchases on yields is negative regardless of investor type, and ($b$) that the reverse effect of higher yields on UST flows is positive. If $cov(a) < 0$ and $cov(b) > 0$, an estimate that confounds ($a$) and ($b$) will be less negative than the true causal effect ($a$), which we wish to estimate. Thus, under our assumptions, simultaneity should lead to an understatement of the impact of FO flows on U.S. yields.

More formally, consider the following example:

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

where $a < 0$ and $b > 0$. It is easy to see that, in the presence of simultaneity, the bias from estimating $a$ by OLS will be in the direction of $b\sigma_1^2/(1 - ba)$, where $\sigma_1^2$ 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 in which FO demand is fully inelastic ($b = 0$) but there exists other price-elastic market segments:

$$y_t = aFO_t + cPR_t + e_3, \text{ and } PR_t = dy_t + e_4, \tag{3}$$

where $PR_t$ are *private* flows into USTs. Here, all flows negatively impact the yield, $a < 0$ and $c < 0$, but private flows are elastic, $d > 0$. The two-equation system above can be re-written as:

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

resulting in an estimate for the price impact of $FO_t$ 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, as long as there exists elastic demand from other market segments.

### 3.2.2 Omitted Variables

U.S. yields are determined by a variety of economic and financial forces. As we discussed earlier, researchers typically attempt to control for the main determinants by including expectations of U.S. real GDP growth and inflation, proxies for financial risk appetite, and fiscal policy indicators. On the other hand, FO demand for USTs is typically assumed to be driven by underlying precautionary, mercantilist, and exchange rate smoothing motives [Obstfeld et al., 2010; Jeanne and Ranciere, 2011; Dominguez et al., 2012] that may depend on the state of the global economy. In turn, the state of the global economy also affects U.S. yields. As we show in Section A2 of the Appendix, global output growth co-moves strongly with non-U.S. advanced economy government yields, which in turn co-vary strongly with FO flows and U.S. yields. Foreign government bond yields, therefore, can be an important source of omitted variable bias in the benchmark model typically used in the literature.

To illustrate the point, suppose U.S. yields depend linearly on domestic factors, FO flows, and also additional foreign factors that are omitted from the analysis:

$$y_t = a_1 X_t + a_2 FO_t + a_3 Z_t + e_t, \tag{5}$$

where $X_t$ represents domestic factors, $FO_t$ is our FO flow variable, $Z_t$ represents other omitted factors, and $e_t$ is an i.i.d. error term. Importantly, suppose $Z_t$ is correlated with both $y_t$, and $FO_t$, which we show is the case in Section A2 of the Appendix for foreign government bond yields.

If Equation 5 is estimated omitting $Z_t$, then the sign of the bias in the estimate of $a_2$ can go in either direction depending on the covariance between the omitted variable and FO flows, $cov(Z_t, FO_t)$, and the covariance between the omitted variable and U.S. yields, $cov(Z_t, y_t)$. Intuitively, omitted variable bias will lead to a less negative estimate of the effect of FO flows on U.S. yields when $Z_t$ covaries positively with both $FO_t$ and $y_t$, or when $Z_t$ covaries negatively with both $FO_t$ and $y_t$. For instance, if $Z_t$ were foreign yield factors that are positively associated with global growth, FO demand for USTs, and U.S. yields – i.e. $cov(Z_t, FO_t) > 0$ and $cov(Z_t, y_t) > 0$ – omitting $Z_t$ will lead to understating the impact of FO flows on U.S. yields, as we find in Table 2. Alternatively, consider the case in which $Z_t$ were Fed LSAP shocks and assume as it is plausible that Fed purchases push down Treasury yields and this occurs during cyclical downturns, when foreign reserve managers more likely sell FX reserves, i.e. $cov(Z_t, y_t) < 0$ and $cov(Z_t, FO_t) < 0$. Omitting such a factor would again lead the researcher to estimate a less negative impact of FO flows on U.S. yields than the true effect.

In practice, the true data generating process is likely much more complex than assumed here. Moreover, it is possible that some factors introduce both omitted variable and simultaneity bias at the same time. For instance, Fed asset purchases can be both counter-cyclical and also introduce simultaneity if the Fed purchases of Treasuries coincide with FO sales. For this reason, the discussion in this section is meant to provide a justification for controlling for these two important factors and an interpretation of our empirical findings, without aiming at a general statement on the sign of this bias. In the rest of paper, therefore, we will focus on providing impact estimates that are robust to endogeneity regardless of the underlying source.

# 4 Identifying a FO Flow Shock via Heteroskedasticity

The regression analysis in the previous section does not control for the possible simultaneity between U.S. yields and FO purchases. To address this thorny identification issue, we estimate a structural VAR (SVAR) for U.S. yields of different maturities and FO flows identified through heteroskedasticity. After describing the VAR specification, we spell out the identification conditions and provide evidence that supports them in this section. We report and discuss our estimation results in Section 5.

## 4.1 VAR Specification

The reduced form VAR that we employ is specified as follows:

$$\mathbf{Y}_t = \boldsymbol{\beta}'\mathbf{Y}_{t-l} + \boldsymbol{\Gamma}'\mathbf{X}_t + \mathbf{u}_t, \tag{6}$$

where

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

and

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

Here, the vector of endogenous variables, $\mathbf{Y}_t$, includes FO flows, 3-month, 2-year, 5-year, 10-year, and 30-year U.S. yields. The yields enter as spreads relative to the Federal Funds Rate ($FF$), which detrends all level variables and also accounts for the systematic component of Fed policy. This setup also permits examining the effects of FO purchases on the U.S. term spread, which is the variable at the center of the Greenspan 'Conundrum' debate.

The vector of control variables, $\mathbf{X}_t$, includes the following domestic factors also used in the specification in Equation (1): expected 1-year ahead real GDP growth, expected 1-year and 10-year ahead inflation, the VIX index, and the structural budget surplus. In addition, to address the concerns with omitted variables discussed earlier, we include foreign yields and U.S. monetary policy shocks. Foreign yields enter as common factors extracted from a panel of short-term and long-term government bond yields for 19 non-U.S. advanced economies, as described in more detail in Section A2 of the Appendix. The U.S. monetary policy shocks are from Swanson [2021], representing innovations to the Fed Funds Target, Forward Guidance, and LSAPs. Finally, the model also includes a small set of dummy variables, represented by $D_t$, to remove three large residual outliers that give rise to wide error bands in the bootstrap procedure used to construct the confidence sets 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.

The VAR above is estimated with 4 lags similar to Bernanke et al. [2004]. In unreported robustness checks, we find that the results are robust to both decreasing and increasing the lag length.

## 4.2 Identifying a FO Flow Shock via Heteroskedasticity

We address the possible simultaneity problem in the relationship between FO flows and U.S. yields by identifying a FO flow shock in the VAR above. To do so, we exploit a shift in the variance of the FO flows and U.S. yields after the Lehman crash in September 2008.

Identification via heteroskedasticity was initially proposed by Rigobon [2003]. Brunnermeier et al. [2021] apply this strategy in a VAR setting. Identification solely based on short-run zero restrictions is difficult to justify in our setting in which asset prices (yields) and quantities (FO flows) are simultaneously determined. Identifying a FO flow shock using sign restrictions is an alternative, but there are other shocks that can generate a negative correlation between FO purchases of USTs and U.S. yields. An example is a monetary policy shock that impacts the Dollar, thereby inducing foreign central banks to sell reserves to stabilize their currency. Finally, relying on external instruments requires sources of exogenous variation in foreign official flows that are hard to isolate credibly in the data, in part because reliable data on central bank reserve holdings and central bank foreign exchange intervention is not available for a sufficiently large cross section of countries.

The basic idea of identification via heteroskedasticity is as follows.[7] Suppose the covariance matrix of the reduced-form residuals, $\mathbf{u}_t$, has at least one break; for example, differing

[7]See Lewis [2022] and Montiel Olea et al. [2022] for a more formal exposition and a short survey of the recent theoretical literature on identification from higher moments, including via heteroskedasticity.

before and after September 2008 as assumed in our application:

$$E(\mathbf{u}_t\mathbf{u}_t') = \begin{cases} \Sigma_1, \text{ for } t = 1, ..., t_{Sep2008} - 1 \\ \Sigma_2, \text{ for } t = t_{Sep2008}, ..., T, \end{cases} \tag{7}$$

so that $\Sigma_1 \neq \Sigma_2$. Further assuming that only the variances change overtime, the two covariance matrices can be expressed as $\Sigma_1 = BB'$ and $\Sigma_2 = B\Lambda B'$, where $\Lambda$ is a diagonal matrix whose elements are the eigenvalues of the matrix $\Sigma_1^{-1}\Sigma_2$ with strictly positive elements $\lambda_{kk}$ ($k = 1, ..., 6$ in our application), and $B$ is the 'impact matrix' to be identified that has $K^2$ unknowns. By adding a second variance regime, the number of identifying equations doubles from $(K^2 + K)/2$ to $(K^2 + K)$, but only $K$ new unknowns are added to the previous $K^2$ unkowns, thus potentially rendering the $B$ just identified. The structural shocks are then recovered from the relationship between the covariance matrix of the reduced form residual ($\mathbf{u}_t$) and that of the structural shocks ($\zeta_t$), which is $\Sigma_{\mathbf{u}} = B\Sigma_\zeta B'$.

More formally, after normalizing the variances in the first regime to one, the matrix $\Lambda$ characterizes the relative change in variances in the second regime. To identify the columns of $B$, all elements of $\lambda_{kk}$ must be distinct [Lütkepohl et al., 2021]. If they are, the impact matrix $B$ is point-identified up to permutations and sign reversals of its columns, the vector of structural shocks $\zeta_t$ can be recovered from the vector of reduced-form residuals $\mathbf{u}_t$, and one element $\lambda_{kk}$ can be normalized to 1 [Lanne and Lütkepohl, 2008]. Labeling and interpreting the shocks of interest can then be accomplished by using additional criteria and information, as for instance discussed by Lewis [2022]. In addition, a critical auxiliary assumption for credibly pursuing this identification strategy is also that all VAR model parameters except the variances are stable over time [Brunnermeier et al., 2021]. It follows that, in this context, the two main threats to credible identification (the main sources of weak identification) are (i) variance changes that are not marked enough across regimes, and (ii) marked variance changes across regimes that are not distinct enough across variables. An example of the latter problem is a variance shift by the same proportion across all variables, akin to a common factor driving shifts in volatilities. As we will illustrate, in our application, the evidence reported strongly supports the conditions needed for point identification.

An appealing aspect of this identification strategy is that we can investigate empirically both sets of assumptions using test statistics and other diagnostic tools on the reduced form VAR parameters and residuals. In the rest of this section, therefore, we review several pieces of informal and formal evidence supporting the critical hypotheses needed for identification of a FO flow shock via heteroskedasticity. We then label and interpret our FO flow shock in Section 5, where we also present the main estimation results on the impact of FO flow shocks on U.S. yields.

Before proceeding, recall here that, as we saw in Section 3.1, the OLS estimate of the impact of FO flows on U.S. yields is substantially stronger in the sample ending in 2007, before the 2008 Global Financial Crisis, compared to the full sample that ends in 2018. This result is consistent with an increase in FO flows volatility after the GFC, although is not conclusive evidence as it could also reflect a covariance change.

## 4.3 Volatility Shifts in FO Flows and U.S. Yields

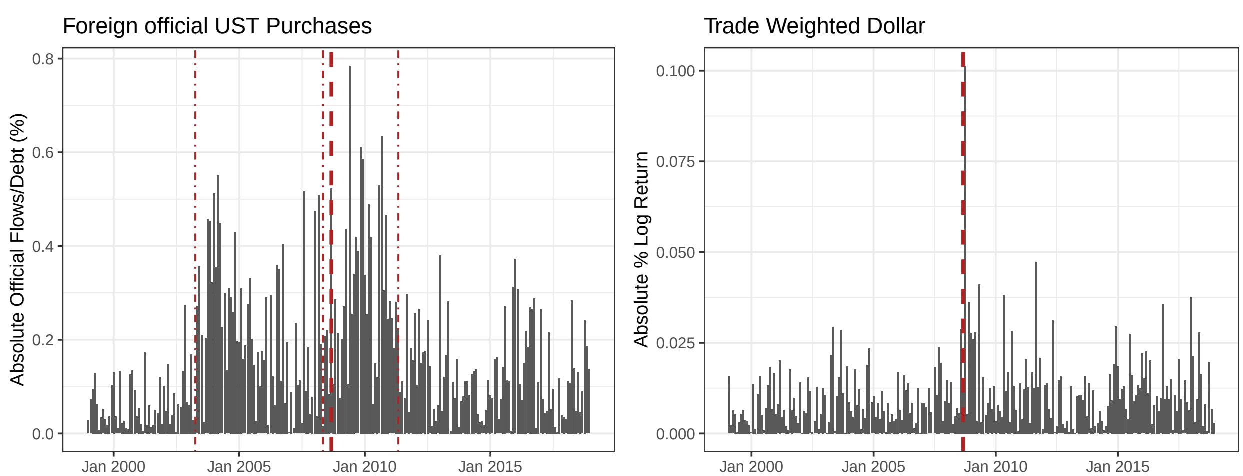
Our critical identification assumption is the presence of a structural break in the variance of the FO flows and the U.S. yields that is not proportional across the variables in the VAR system.[8] So, we will now review the evidence on the following: (i) structural breaks in capital flow patterns and the U.S. Treasury market documented in the existing literature; (ii) statistical test results that assume a *known* break point in variances; (iii) tests that assume *unknown* break points; and (iv) evidence ruling out proportional variance changes.

[8] Lewis [2021] and Montiel Olea et al. [2022] note that SVAR identification can also be achieved under the weaker conditions. For example, identification could be achieved if there is time-varying volatility in the data and conditional independence of the structural shocks. Strong evidence of residual non-normality and heteroskedasticity is reported in Appendix A3. As the evidence in favor of a non-proportional structural break in our data is compelling, we do not need to invoke these weaker assumptions.

### 4.3.1 Evidence from Other Studies

Studies exploiting heteroskedastic-based identification often draw their credibility from external knowledge of historical events. In our case, we assume a structural break in September 2008 in coincidence with the Lehman crash also because of a shift in the pattern of international capital flows before and after the GFC along multiple dimensions that is well-documented in the literature. For example, Ahmed and Zlate [2014] find increased sensitivity of flows to interest rate differentials, unconventional monetary policies, and capital controls after 2008. López and Stracca [2021] document that capital flows fell abruptly after the 2008 crisis and that their composition shifted from bank to portfolio flows. Forbes and Warnock [2021] shows that drivers of capital flows changed from risk appetite to commodity prices Erik et al. [2020] show that the U.S. Dollar has gained importance since the 2008 crisis as a risk factor. Recent evidence from Du et al. [2022] also documents a regime change in U.S. Treasury markets following the GFC, coinciding with the drastic drop in U.S. yield volatility as the Fed Funds Rate approached the zero lower bound.

Figure 3: Absolute FO Flows and U.S. Dollar Returns




The figure plots absolute values of FO flows and monthly log-returns of the trade-weighted U.S. Dollar in the left and right panels, respectively. The thick dashed vertical line is September 2008. The thin dot-dashed vertical lines are detected breaks in April 2003, May 2008, and May 2011 in the absolute FO flows series using the testing framework in Bai and Perron [2003]. For absolute FO flows (absolute Dollar returns), the pre September 2008 mean is 0.15 (0.009) and the post September 2008 mean is 0.18 (0.127).

The left panel of Figure 3 plots the absolute value of FO flows, which is a measure of volatility, while the right panel plots the absolute value of the monthly log-returns of the trade-weighted U.S. Dollar. According to this measure, the volatility of both FO flows and the U.S. Dollar returns, on average, increased in the post-2008 period. In fact, the pre-September 2008 mean absolute FO flow and absolute Dollar return are 0.15 and 0.009, respectively, while the the post-September 2008 means are 0.18 and 0.127. Appendix A3 also reports strong evidence of reduced-form VAR residual non-normality and heteroskedasticity consistent with the assumptions made for identification via heteroskedasticiy.

#### 4.3.2 Known Break Point Tests

In line with the structural shifts in capital flows documented by other studies, we also find formal statistical evidence that the volatility of FO flows changed since 2008. Indeed, the standard deviation FO 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). These large and statistically significant changes in FO flow volatility are robust to assuming alternative break dates around the GFC or excluding the GFC crisis period altogether. In unreported analysis, we also find similar changes in FO flow volatility when inspecting large country holdings of USTs as in the case of China. Finally, the annualized volatility of the U.S. Dollar also increased, practically doubling, from 3.77% to 6.14% in the period after September 2008. Again, the change in variance is significant at the 1% level.

Table 3: VARIABLE VOLATILITIES BEFORE AND AFTER SEPTEMBER 2008

| | $FO_t$ | $y_{us,t}^{3M-FF}$ | $y_{us,t}^{2Y-FF}$ | $y_{us,t}^{5Y-FF}$ | $y_{us,t}^{10Y-FF}$ | $y_{us,t}^{30Y-FF}$ |
|---|---|---|---|---|---|---|
| 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*** |

The table report the sample variances within each period for the endogenous variables in Equation 6. In the F-test, the null hypothesis is that the ratio of variances equal 1.

In addition to a shift in FO flows volatility, Table 3 reports the variance changes in the U.S. yields that enter the VAR above as endogenous variables. Recall here that for identification purposes, all but one variable must undergo a variance shift of different proportion. Associated with the Fed Funds 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, but by different degrees across maturities. Post-2008, U.S. yield spread variances are 60-90% of the pre-September 2008 period variances. An increasing variance of FO flows and a non-proportional decreasing variance of all remaining variables are in line with the necessary conditions for point identification as discussed earlier and in Lewis [2021] and Lütkepohl et al. [2021].

#### 4.3.3 Unknown Break Point Tests

Next, we focus on tests for multiple volatility changes when the break points are unknown, using the procedure in Bai and Perron [2003]. As a monthly proxy for FO flow volatility, we use the absolute value of FO flows plotted in Figure 3, considering also squared foreign official flows for robustness.

The Bai and Perron [2003] test detects three breaks in this series based on the Bayesian Information Criterion: April 2003, May 2008, and May 2011. May 2008 and May 2011 coincide with the onset of the GFC in the United States and the beginning of the European Crisis, respectively. The 90% confidence interval for the May 2008 break point, in particular, contains the assumed break in September 2008. On the other hand, April 2003 lines up well with the resumption of global capital flows after the emerging markets crises of the late 90's and early 2000's and the burst of the dot com bubble in 2001 in the United States. The April 2003 break also aligns well with the beginning of the foreign exchange reserves accumulation process in emerging markets that followed the preceding decade of financial crises. Similar break dates are identified in the time-series of squared FO flows: April 2003, December 2007, and December 2013 (with September 2008 again contained within the confidence bands).

While in principle the two additional beak dates could be exploited as a source of overidentification restrictions (but only if variance shifts in the U.S. yields also occurred on these dates), in our application, we only consider the 2008 break date, which lines up well with the break in the variances of the yields. This is to helps us keeping the statistical analysis of the VAR system manageable, since the identification conditions become more complex and difficult to test statistically with 6 variable [Lütkepohl et al., 2021]. Like Lanne and Lütkepohl [2008], therefore, we prefer to specify our baseline model with fewer regimes than those identified via statistical tests. We also note that structural estimation remains consistent under mispecified break dates [Rigobon, 2003; Sims, 2020].

### 4.3.4 Ruling out Proportional Variance Shifts

As we noted earlier, credible identification via heteroskedasticity is also threatened by the possibility that all variables in the VAR change by the same proportion.[9] If all variances change by the same proportion, there is no new identifying information as the covariance matrices become scalar multiples of each other across regimes. Referring again to Table 3, it is quite unlikely that all variables experience a volatility shift of the same proportion in our application given that the variance of the FO flows increases, while the variance in the U.S. yields declines over time by seemingly different degrees across maturities. Nonetheless, following Lütkepohl et al. [2021], we formally test the VAR residuals to exclude this possibility by using the joint Wald test for the null hypothesis that the variance changes are proportional:

$$H_0 : \lambda_1 = \lambda_2 = ... = \lambda_6. \tag{8}$$

The Wald statistic for this test is 55.70 with 20 degrees of freedom clearly rejecting the null of equal variance shifts and suggesting that there is at least one distinct regime shift in volatility across our VAR variables.

[9] An example, in our context, is a common shock to the volatilities of all variables in the VAR system, such as the Great Moderation, or the increase in global volatility during the 2008 GFC.

Lütkepohl et al. [2021] proposes a sequential approach to test all possible permutations in (8) but this test has sharply decreasing power as the number of variables and lags in the VAR increases. Taking this strong limitation into consideration, we supplement the test on the joint equality of all permutations of the elements of $\Lambda$ with a battery of tests on the equality of selected permutations of the $\lambda_k$ elements detailed below. Specifically, we implement this test on progressively smaller VARs, while always including the FO flows variable. Thus, we sequentially remove the following variables: the 3-month yield; the 3-month yield and 2-year yield; the 3-month, 2-year, and 5-year yields; and the 3-month, 2-year, 5-year, and 10-year yields. The procedure leads us to analyze 5, 4, 3, and 2-variable VAR models, respectively. The null hypothesis, which is always the same as in (8), is rejected for all smaller VARs with test statistics of $58.53^{***}$, $9.22^{**}$, $9.52^{*}$, and $8.051^{**}$, respectively).[10] Finally, we consider the following set of bivariate VARs: FO flows and the 2-year yield; FO flows and the 5-year yield; and FO flows and the 10-year yield, finding that the null hypothesis of variance equality is again always rejected with test statistics of $27.34^{***}$, $11.621^{***}$, $8.58^{**}$, respectively.

## 4.4 VAR Parameter Stability

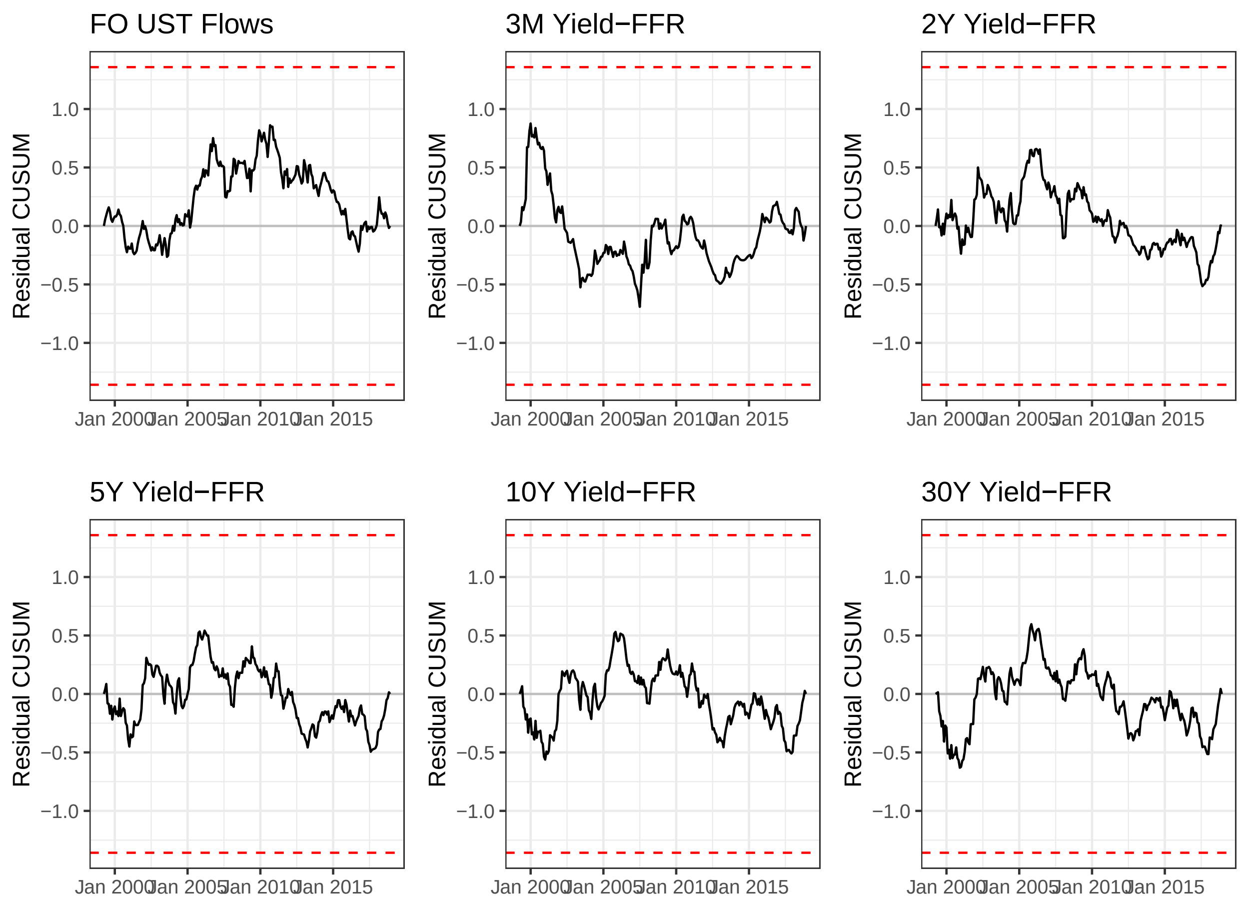
An important auxiliary assumption for credibly pursuing identification via heteroskedasticity is also that all VAR model parameters and the structural covariance terms are stable over time, and just the variances of the shocks change. Importantly we find no evidence of VAR parameter instability, consistent with the findings in Sims and Zha [2006] in favor of VAR specifications with time-varying volatility rather than time-varying coefficients.

Whether or not the 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 4 reports these CUSUM plots equation-by-equation and shows that there is no evidence of instability or breaks in the VAR parameters since all series are well within the 95% confidence band

[10] '***', '**', '*' denote significance at the 1, 5, and 10 percent levels, respectively.

region.

Figure 4: CUSUM Test for VAR Parameter Stability




Solid lines indicate cumulative standardized VAR residual series from Equation 6. 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.

# 5 SVAR Estimation Results

Having established that the data support the statistical assumptions needed for valid identification via heteroskedasticity, in this section, we discuss the labeling and interpretation of our FO flow shock. We then report our main estimation results on the impact of such shock on U.S. government yields.

We estimate the reduced-form VAR by OLS with benchmark-consistent FO flow data and then recover the structural shocks following the procedure outlined in Lanne and Lütkepohl [2008]. In this procedure, the matrices $B$ and $\Lambda$ corresponding to the covariance matrices in (7) are initialized by maximizing the log-likelihood under the assumption of Gaussian residuals $\mathbf{u}_t$. The VAR parameters, $B$ and $\Lambda$, are then estimated iteratively using generalized least squares. The structural shocks can be recovered using the estimated $B$ matrix and the reduced-form residuals from $\zeta_t = B^{-1}\mathbf{u}_t$.

## 5.1 Labeling and Interpreting a FO Flow Shock

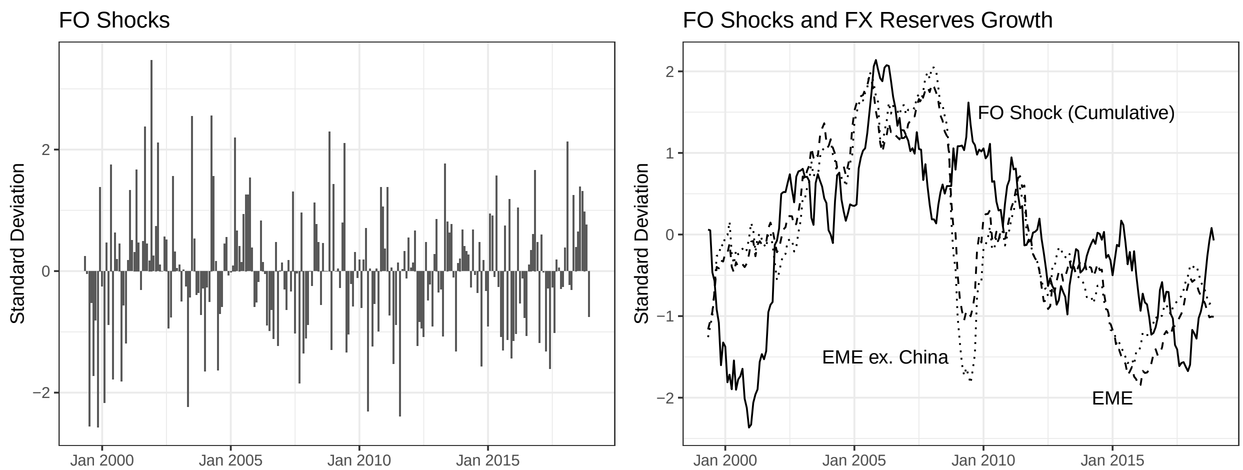
There is no set technique for labeling and interpreting the shocks identified via heteroskedasticity. Lewis [2021] and Brunnermeier et al. [2021] provide a number of alternative suggestions in light of which we follow a multi-pronged strategy. The strategy includes (i) selecting the shock that produces the largest time-zero responses in the FO flow variable (also proposed by Brunnermeier et al. [2021]), (ii) checking that the impulse responses are theoretically consistent with the negative covariance arising from the causal effect of flows on yields, (iii) 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 (iv) directly inspecting the elements of the $B$ matrix. The estimated $B$ matrix is:

$$\hat{B} = \begin{bmatrix} -0.079 & 0.022 & -0.005 & 0.083 & -0.023 & -0.080 \\ 0.043 & 0.110 & 0.073 & 0.017 & 0.044 & 0.069 \\ -0.017 & 0.008 & 0.235 & 0.004 & -0.036 & 0.066 \\ -0.055 & 0.027 & 0.231 & -0.028 & -0.077 & 0.113 \\ -0.093 & 0.032 & 0.199 & -0.050 & -0.051 & 0.128 \\ -0.109 & 0.012 & 0.131 & -0.039 & -0.023 & 0.163 \end{bmatrix}. \tag{9}$$

Column 6 of the $B$ matrix represents our FO flow shock. This column has the largest negative loading in the first element (corresponding to foreign official flows) but positive elements on all yields. The shock corresponding to column 6 is also the only one that satisfies all other criteria listed above. This shock explains roughly 30% of the FO flow series FEVD.

The left-panel of Figure 5 plots our estimated FO flow shock and the right-panel cumulates the shock series. The cumulated FO shock series closely comoves with measures of foreign exchange reserve growth for emerging markets. The correlation between the cumulative FO shock and the year-on-year growth in official reserves is 0.58 including China and 0.48 excluding this country, both statistically significant at the 1% level.

Figure 5: Identified FO Flow Shocks and Emerging Market Reserves




The left panel plots the standardized structural FO flow shocks identified via heteroskedasticity from the estimated VAR residuals in Equation 6. The right-panel plots the cumulated FO shocks (solid line) against year-over-year growth of foreign exchange reserves (excluding gold) of emerging markets (dashed line) and the reserves growth of emerging markets excluding China (dotted line).

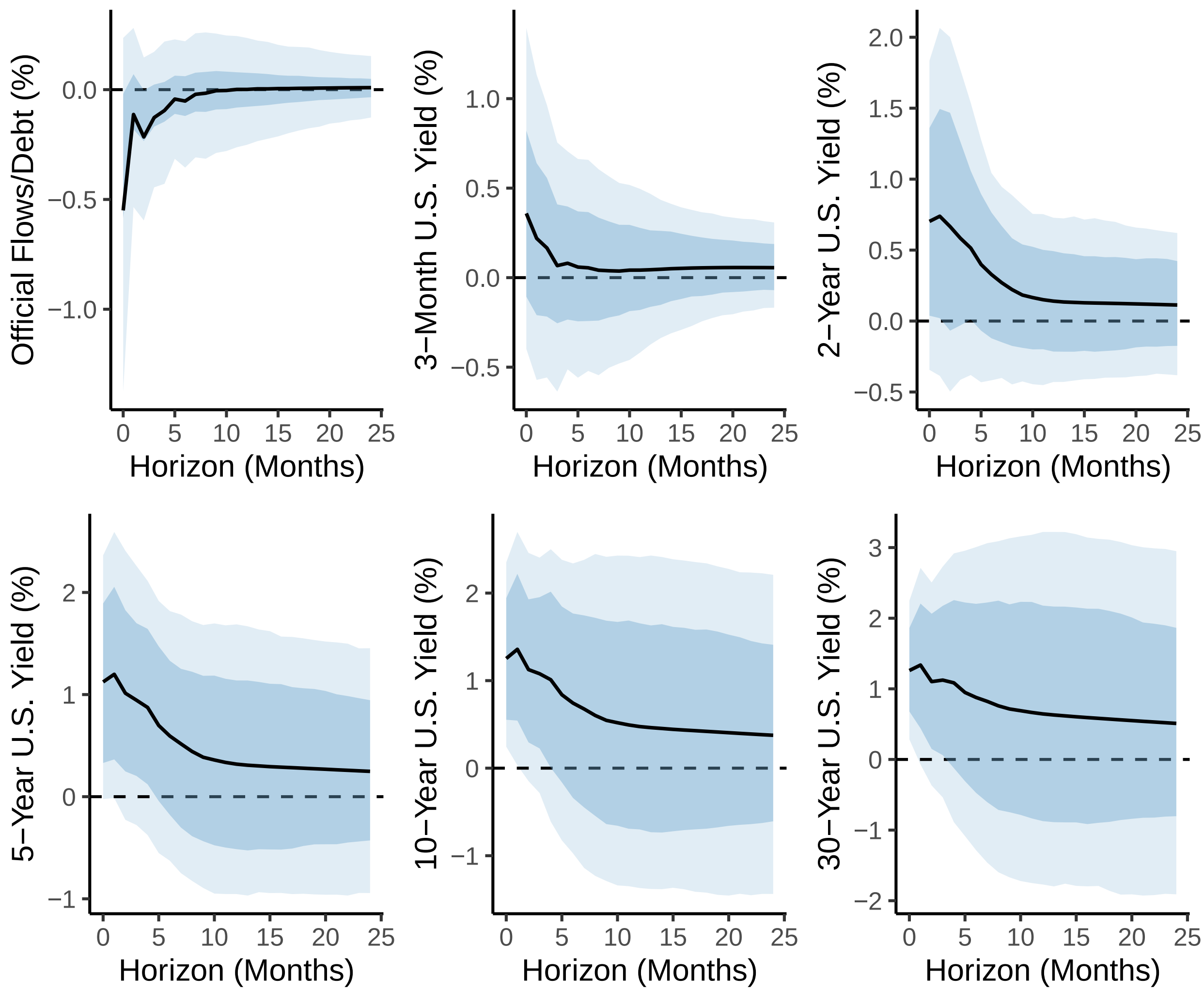
We are agnostic on the interpretation of the other five structural shocks in our VAR system. Nonetheless, the estimated VAR is assumed to have a fundamental and inevitable representation. Conceivably, one could therefore assume that the other yield differentials carry information on the level, the slope, and the curvature of the U.S. yield curve, as well as information over fundamental objectives of monetary policy such as the inflation and the state of the labor market as captured by the output gap.

## 5.2 FO Flow Shocks and U.S. Yields

Figure 6 reports the main results of the paper. It plots the boostrapped mean impulse response functions (IRFs) to an identified FO *sale* shock, together with confidence bands for all variables in the VAR. As the VAR is a linear model, the IRFs to a FO *purchase* shock are the same, but with the opposite sign. 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 one of the two regimes to back out the structural shocks from the reduced-form residuals, Figure 6 plots IRFs rescaled to reflect a $100 billion foreign official sale (or 0.55% of marketable debt based on December 2017 debt outstanding) as the benchmark estimates in Table 2 of Section 3.

The FO flows' own response shows that the shock dies out within 5-6 months (the upper-left panel plot). Following a $100 billion FO sale, there is no statistically significant impact on the 3-month yield. The 2-year yield rises about 70 basis points on impact, but the effect is only significant at the 68% level, consistent with the notion that shorter-dated U.S. yields are largely determined by monetary policy, while foreign purchases affect the longer end of the U.S. term structure, via the term premium. The 5, 10, and 30-year yields rise by 112, 125, and 126 basis points on impact, respectively, and are statistically significant at the 90% level. The responses peak within a month, halve within about half a year in line with the persistence of the identified FO shock, and then decay more slowly thereafter. At the one-year horizon, the impact of the shock on the 10-year and the 30-year yields is still about 50 bps, close to the upper end of the range of effects estimated in the extant literature. Because our analysis studies the spread between yields and the Fed Funds Rate, the results further suggest that FO purchases and sales of USTs can drive a significant and persistent wedge in the pass-through of policy rates to long-term Treasury yields, consistent with Greenspans's 'Conundrum' observation and the 'Global Savings Glut' Hypothesis, even at a horizon of one

Figure 6: Impulse Response of U.S. Yields to a FO Sale Shock
(In Percent)




The figure plots IRFs of U.S. yields as spreads over the Fed Fund Rate to a negative FO sale shock (i.e., a net sale, top-left panel). The reduced form model is a VAR(4) specified as in Equation 6 and includes FO flows scaled by marketable debt lagged 12 months. Light and dark shaded regions represent 90% and 68% bootstrapped confidence intervals, respectively, based on 1,000 bootstrapped samples. The solid line is the bootstrap mean IRF estimate. Shocks are scaled to reflect \$100B a FO sale of USTs.

year or longer. Finally, a FO sale shock does leave the U.S. term spread persistently steeper as we can see by taking the 10-year IRF and subtracting the 3-month IRF. The wide error bands, however, suggest that these longer-term effects not estimated precisely in our model.

## 5.3 Economic Significance

The short-term impact of our FO flow shock is much larger than the 13-68 basis point range typically found in the literature, or the 19-44 basis point benchmark estimate in Table 2 of Section 3, estimated by assuming that FO flows are inelastic and without controlling for U.S. monetary policy shocks and foreign yield dynamics. In this section, therefore, we elaborate on their economic significance, considering the March 2020 episode of distress in the U.S. Treasury market and a hypothetical small change in the composition of foreign reserves, following the imposition of sanctions on Russia official reserves in 2022. Our estimated contemporaneous price impacts are large also relative to the literature on quantitative easing (i.e. domestic official purchases of USTs), a comparison on which we focus at the end of the section.

### 5.3.1 Price Impact of FO Sales of USTs in March 2020

Amidst the onset of the COVID-19 pandemic that disrupted global financial markets, briefly also disrupting the U.S. Treasury market, in March 2020, foreign investors sold roughly \$300B in USTs, approximately \$150-200B of which have been tied to sales by foreign official institutions [Vissing-Jorgensen, 2021; Weiss, 2022]. Liquidity provision in U.S. dollars to the local interbank markets and direct foreign exchange interventions are both considered important drivers of these foreign official sales. During this episode, domestic private holders such as hedge funds and mutual funds also sold USTs for liquidity reasons, while the Federal Reserve engaged in a new round of QE to stave off the crisis.

For context, at the peak of financial stress from March 9 to 18, the 10-year nominal yields rose 64 basis points while real yields (default-adjusted) rose 100 basis points. Treasury yields started falling on March 19, after the first announcement of Treasury purchases by the Federal Reserve on the preceding day. Because the bulk of the financial distress occurred within this 10-day period but we only have monthly frequency measures of FO purchases, we assume that one-third (or 10 days out of a month worth of flows) of the total $150-200 billion sales occurred over the 10-day stress period of March 9-18. Based on our identified impacts in Figure 6 and assuming a FO sale range of $50-$67 billion over this critical 10-day period, we estimate a 56-80 basis point impact on the U.S. 10-year yield caused by FO sales, which is quite sizable but not inconsistent with the total changes in the nominal yield noted above because it abstracts from sales or purchases by other investors and the Federal Reserve.

#### 5.3.2 A Shift away from Dollar Reserves by China and Saudi Arabia

We now consider the implied price impact of a hypothetical reduction in central bank holdings of USD-denominated assets by 1 percent. We consider China and Saudi Arabia, the major official holder of U.S. Dollar assets and a medium-size holder like Brazil or South Korea. Over the last two decades, China has operated a fixed exchange rate regime first and then a managed float, accumulating a massive war-chest of official reserves, with total official reserves (excluding gold) of $3.25 trillion as of March 2022. Saudi Arabia has long operated a hard peg to the Dollar and is the largest oil exporter. It official reserve holdings are estimated at $326 billion in March 2022. After the imposition of sanctions on Russia because of the invasion of Ukraine, official reserve holdings have lost some of their precautionary value. Going forward, it is therefore plausible to anticipate a composition change in favor of physical gold holdings, strategic reserves of energy and other scarce commodities, and perhaps crypto assets, or at least a slowdown in the pace of accumulation of Dollar assets.

Table 4 reports a back-of-the-envelope estimate of the impact on U.S. yields if either of these two countries were to shift 1% of their reserves away from USTs. The calculation relies on a set of simplifying assumptions. First, we assume that all FO flows are driven by reserves accumulation and management, while in fact they also include other government entities' transactions such as sovereign wealth funds. Second, we assume that all USD-denominated assets are in the form of Treasury securities. Third, as customary, we assume that the Dollar share in total reserves is 60% in line with the estimates reported by the IMF Currency Composition of Foreign Exchange Reserves (COFER).[11]

Table 4: Impact of a 1% Reduction in Dollar Reserves

| | China | Saudi Arabia |
|---|---|---|
| FX Reserves ($Bl, Mar 2022) | 3,250 | 326 |
| Assumed %USD Share | 60% | 60% |
| 1% Outflow ($Bl) | -19.5 | -2 |
| 5Y yield elasticity per $1Bl | 1.12bps | 1.12bps |
| **Contemporaneous impact on 5Y yield** | **+21.8bps** | **+2.2bps** |
| 10Y yield elasticity per $1Bl | 1.25bps | 1.25bps |
| **Contemporaneous impact on 10Y yield** | **+24.4bps** | **+2.5bps** |

Author's calculations assuming that all USD-denominated reserve assets are U.S. Treasury securities. Calculations further assume a share of USD assets of 60% based on data from the IMF Currency Composition of Official Foreign Exchange Reserves (COFER). Yield elasticities are from the VAR model estimated in Equation 6.

In this scenario, in the case of China, a 1% shift away from the Dollar amounts to a -$19.5 billion outflow ($\$3,250\text{B} \times 0.6 \times -0.01$). Using the estimated contemporaneous impact from Figure 6 above of 1.12 and 1.25 basis points per $100 billion flows, 5-year and 10-year yields could rise by 21.8 and 24.4 basis points, respectively. The same 1 percent reallocation by Saudi Arabia could result in a -$2 billion outflow, implying the 5-year and 10-year yield rise by only 2.2 and 2.5 basis points, respectively. Based on these calculations, we conclude that while a reduction of the Dollar share of reserves in medium-size countries like Saudi would be more easily absorbed by the UST market, if China were to shift its reserves composition away from the Dollar it might have a much stronger impact, comparable in magnitude to the effects of Fed QE.

[11] The exact share of global USD reserves reported by COFER as of Q4 2021 is 58.81%.

## 5.4 Comparison with the Effects of Quantitative Easing

A closely related literature studies the effects of Federal Reserve asset purchase programs, or QE, on U.S. interest rates. Estimates in these studies vary, but they are generally smaller than the estimated impacts of FO purchases discussed above. Broadly speaking, a $100 billion purchase via QE has been linked to a decline in long-term U.S. yields ranging from 5 to 15 basis points [Gagnon et al., 2011; D'Amico and King, 2013; Swanson, 2021; Rebucci et al., 2022], although some studies find larger effects.

There are several reasons why QE and FO purchases or sales of USTs might affect U.S. bond yields differently. First, while domestic asset purchases are swaps of assets of different duration, foreign flows entering the U.S. economy provide incremental resources and purchasing power [Kohn, 2016]. Second, another important issue is omitted variable. Similar to most studies on FO flows and U.S. yields, most QE studies also do not control for global cyclical factors which might influence domestic market conditions. Fed QE purchases may be meeting also foreign investors' sales, possibly leading to a smaller estimated price impact on U.S. yields compared to the same QE purchase amid normal market conditions. Beltran et al. [2013] stress 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 rounds of QEs being priced in bond markets even prior to their announcement. Finally unlike FO purchases, domestic QE programs may be associated with inflation uncertainty, which would put upward pressure on yields.

# 6 Robustness

Before concluding, in this section, we report the results from a number number of robustness checks on the main results. The benchmark is the VAR model with both domestic controls and foreign yields as reported in Figure 6. Here, we consider the following alternative specifications: a VAR model with only domestic controls variables, like in most previous studies, to assess the importance of controlling for the foreign government bond yields; a VAR model that endogenizes foreign *private* flows; a VAR model that endogenizes the long-term and short-term foreign yields rather than simply holding them constant as control variables, where our two foreign yield factors now enter as a spread, $\mathcal{Y}^{10Y}_{g,t} - \mathcal{Y}^{3M}_{g,t}$ to keep the VAR system as small as possible; and, finally, a VAR that endogenizes foreign and private flows, and foreign yield factors.[12] Endogenizing private flows can account for foreign private investors responding to higher yields and possibly absorbing excess demand from FO players. Similarly, U.S. and foreign yields can affect each other as these markets are highly integrated in advanced economies and may both reflect evolving global economic conditions as we also illustrated in Appendix A2.

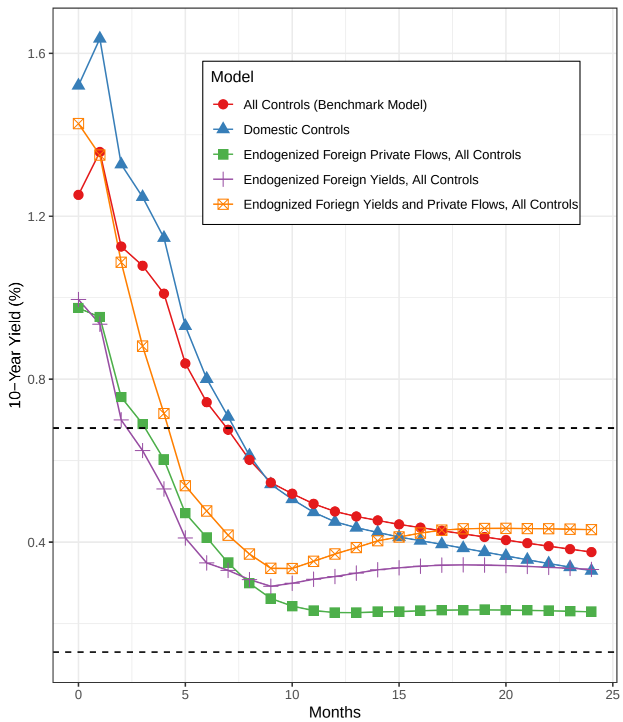
Figure 7 summarizes the main results comparing the same IRF from all the considered model specifications. To conserve space, we focus on the IRF of the 10-year yield. The contemporaneous impact across the five alternative model specifications range from 100 to 150 basis points. All estimates, however, are well outside the 13-68 basis point range typically found in the literature. While all contemporaneous impacts are substantially larger than typically estimated, in all models, the effects decay in line with the persistence of the shock documented earlier, eventually falling within the typical range.

Figure 8 reports the full set of IRFs from the VAR model that endogenizes foreign private flows, foreign yield factors, and also controls for both domestic factors and foreign yields.[13] The results are qualitatively and quantitatively very similar to the baseline VAR in Figure 6. Further validating that a FO flow shock is well identified, the contemporaneous response of the *foreign private* flows is insignificantly different from zero, and slightly positive, implying that the identified FO sale of USTs is not matched by private investor purchases. The response of the common factor in the foreign term spread is also insignificant, consistent with the implicit assumption in our baseline model that an identified FO Treasury purchase

[12] Recall, here, that U.S. monetary policy is controlled for throughout all these experiments.

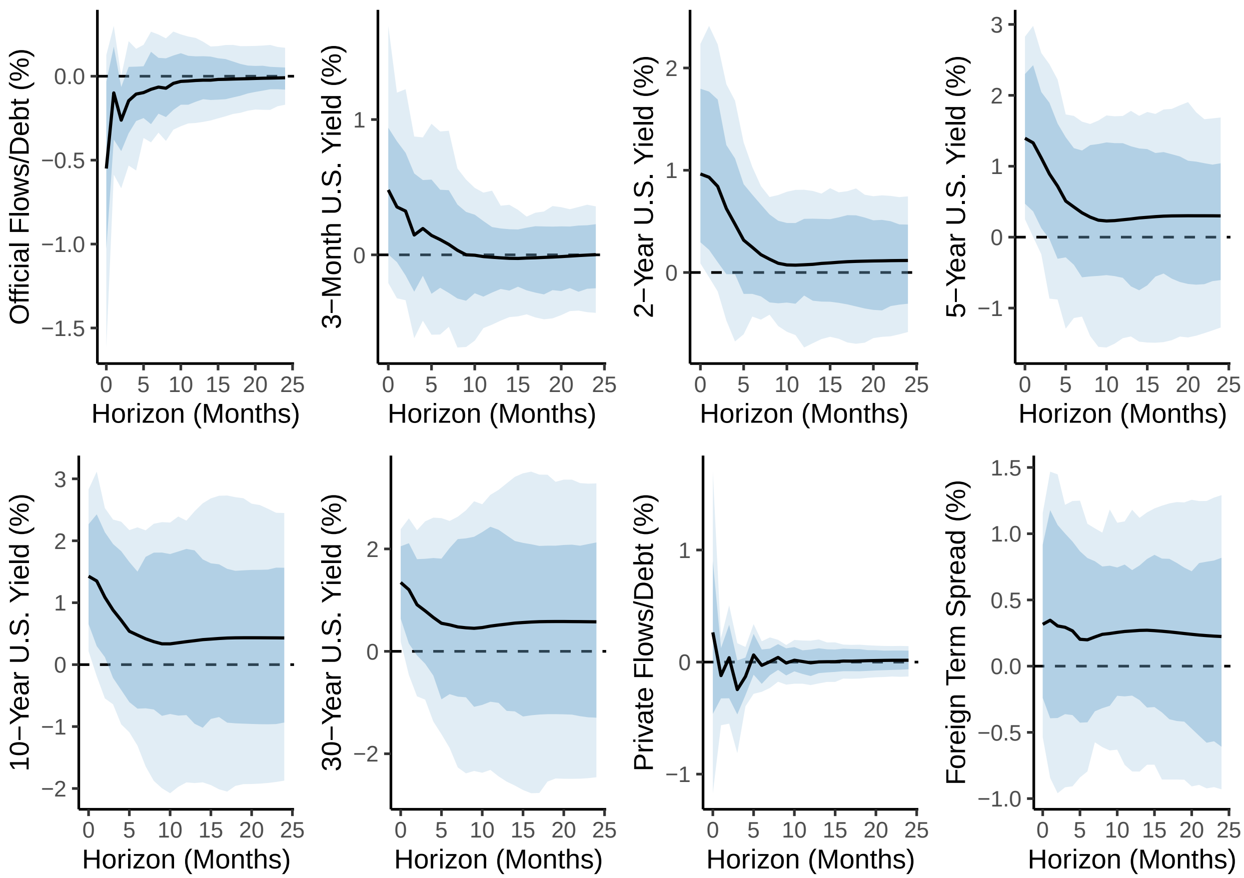
[13] Due to computational constraints, the confidence intervals are based on 100 bootstrapped samples.

Figure 7: Impulse Response of the 10-year Yield Yields to a FO Sale Shock: Alternative SVAR Model Specifications




The figure plots the IRFs of the 10-year yield from alternative specifications of the VAR(4) in Equation 6. The (red) line with circles is the benchmark model in Equation 6. The (blue) line with triangles is a VAR model with only domestic control variables; the (green) line with squares is a VAR that endogenizes foreign *private* flows; the (purple) line with crosses is a VAR that endogenizes the long-term and short-term foreign yields rather than simply holding them constant as control variables; and, finally, the (yellow) line with boxes is a VAR that endogenizes foreign and private flows, and foreign yield factors. Responses are scaled to represent a \$100B foreign official sale of USTs. The dashed horizontal lines indicate the range of estimates from previous studies [-13 bps,-68 bps] in Table 2.

Figure 8: Endogenous Foreign Private Flows and Foreign Yields




This figure plots the impulse responses from the VAR(4) in Equation 6 with foreign private flows and foreign yield factors entered as endogenous variables, and corresponding to the IRF "Endogenized foreign yields and private flows, All controls" model in Figure 7. FO flows are benchmark-consistent scaled by U.S. marketable debt lagged 12 months. Light and dark shaded regions refer to 90% and 68% bootstrapped confidence intervals based on 100 bootstrapped samples. Solid line is the bootstrap mean IRF estimate. Shocks are scaled to reflect $100B foreign official sale of USTs.

or sale shock should disproportionately impact U.S. yields. Note however that the foreign term spread remains positive over time with little or no sign of reversion toward its initial value, matching the persistent increase in the U.S. term spread.

We conclude from this exercise that our analysis is quite robust. Contemporaneous effects of FO flow shocks on U.S. yields are at least 30% larger than the estimates reported in previous studies, converging to the latter within four-six months.

# 7 Conclusions

Exploiting a volatility shift in foreign official flows and U.S. yields in 2008, we identify a FO flow shock in a SVAR via heteroskedasticity. We estimate that an identified $100 billion sale of USTs by foreign officials raises the 10-year yield by more than 100 basis points on impact before falling within the range of values reported in previous studies within a half year. These results bear important implications for the U.S. Treasury market functioning and U.S. financial conditions. Back-of-the-envelope calculations suggest that a 1 percent shift away from Dollar reserves held by the central bank of China could impact the U.S. Treasury market as much as previous round of QE. Similarly, we find that FO sales of USTs have been responsible for the lion's share of the U.S. Treasury yield increase during the March 2020 episode of distress.

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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. Section A2 explains how we construct our foreign yield common factors and describes their cyclical properties, which are relevant to set priors on the omitted variable bias that we document in the paper. Section A3 reports VAR residuals diagnostic statistics that support our identification approach through heteroskedasticity.

## A1 Data Sources and Summary Statistics

Sources and definitions of all variables used in the analysis are collected from a variety of outlets. Monthly long-term and short-term government bond yields used to to construct foreign bond yield factors 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. In addition to U.S. interest rate data, we employ data for 19 other countries: Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, The United Kingdom, Ireland, Italy, Japan, Luxembourg, Netherlands, Norway, New Zealand, and Sweden. The quarterly nominal gross domestic product (GDP) in U.S. dollars is also taken form the OECD for all these countries. Quarterly GDP 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.

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 |

The sample period is from January 1999 to February 2021. Yield, GDP growth, inflation and VIX statistics are in percentages. The structural budget surplus 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) are linearly interpolated to monthly frequency.

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 reading 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 federal 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). Summary statistics for the main variables used in the empirical analysis are reported in Table A.1.

# A2 Foreign Yield Factors and their Cyclical Properties

In this section we describe the construction of the foreign common yield factors that we use in the empirical analysis and discuss their cyclical properties.

## A2.1 Estimating Common Factors in Foreign Yields

We construct both a long-term and short-term yield factor. We use monthly 3-month and 10-year government bond yields for the same 19 advanced economies above, from January

1999 to February 2021. We consider only advanced-economy government bonds as they tend to exhibit lower credit risk than emerging-market bonds [Du and Schreger, 2016]. Including emerging market yields in the construction of the common foreign yield factors would also be inappropriate because, save Japan and Switzerland, most large foreign official UST holders are indeed emerging markets. Unlike emerging market yields, advanced economy yields are more likely to reflect counter-cyclical global investor demand for safe assets, a variable that we wish to control for.[A1]

Several approaches can be taken to construct common factors. The estimated first principal component (PCA) of our 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 at time $t + h$. This might not be desirable in our context, as we model forward looking variables. An alternative approach that doesn't suffer from this drawback is taking the cross-section average (CSA) of yields in each period (see, for example, Cesa-Bianchi et al. [2020]).

Practically speaking, the difference between factor estimates using the PCA approach and the CSA approaches boils down to the weights assigned to each component. The PCA approach estimates weights for each component, while the CSA approach applies equal or economically interpretable weights to each individual country yield. Weights based on economic rationale (e.g. GDP shares or market debt outstanding shares) have the two advantages. First, because the weights are not estimated there is less estimation uncertainty. Second, weights can vary over time.

Weights usually are economically motivated depending on context. We estimate long and short-term bond yield factors as the GDP-weighted averages of long-term and short-term bond yields across countries. GDP weights are intuitive since larger countries, based on economic activity, will be more influential in determining variation in the foreign yield factor. Alternatively, weights reflecting differences in the supply of national public debt

[A1] Del Negro et al. [2019] finds that rising demand for safe assets is an important factor driving global interest rates, as are global trends in economic activity.

across countries could be considered. Specifically, we use lagged U.S. dollar-denominated GDP as follows:

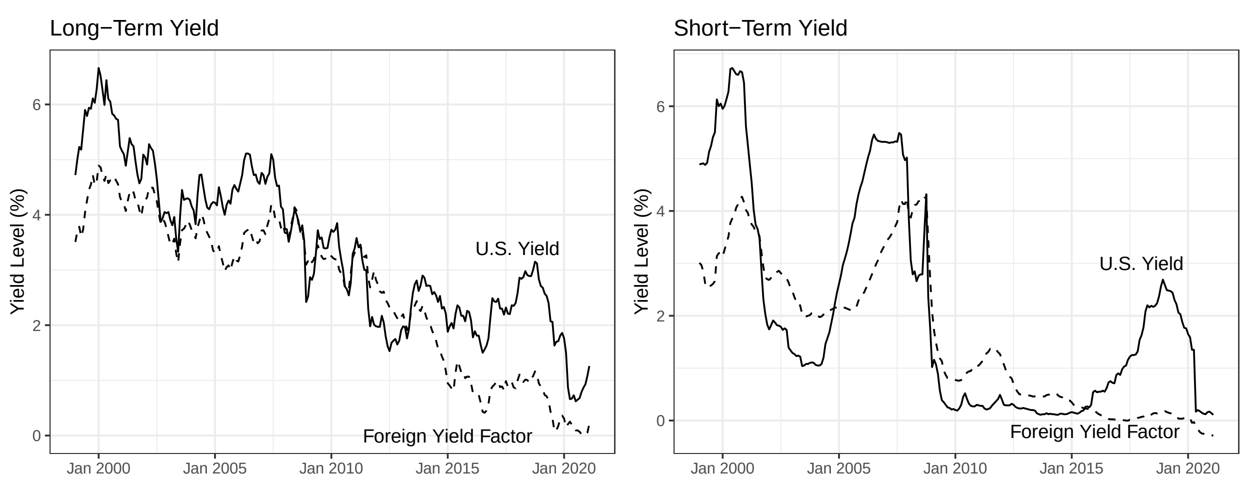
$$\mathcal{Y}_{g,t}^{10Y} = \sum_{i \notin us} w_{i,t-1} y_{i,t}^{10Y}, \quad \mathcal{Y}_{g,t}^{3M} = \sum_{i \notin us} w_{i,t-1} y_{i,t}^{3M}, \tag{A.1}$$

where he 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. Thus, $w_{i,t} = GDP_{i,t}/\sum_{i \notin us} GDP_{i,t}$. Note here that the U.S. yields are omitted because they are the dependent variable of interest in our empirical analysis.

## A2.2 Cyclical Properties

Foreign common yield factors, U.S. yields and Treasury holdings, including official and private foreign demand, are intertwined. Here, we want to illustrate the critical co-movements, which are a potential source of omitted variable bias as discussed in the paper, without establishing any causal relationships among these variables.

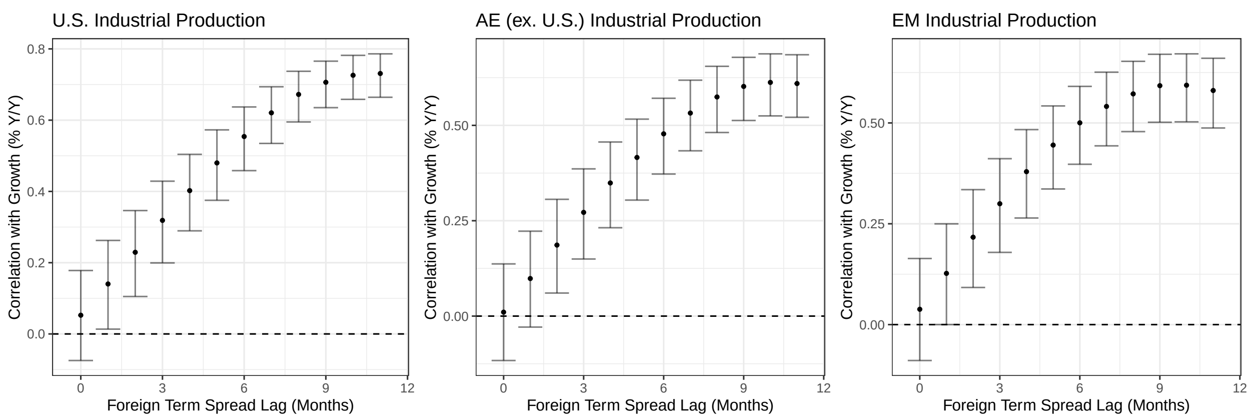
Figure A.1: Common Foreign Yield Factors and U.S. Interest Rates




The solid lines are the U.S. 10-year (left panel) and 3-month (right panel) yields. The dashed lines are the estimated 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.1.

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 10-year yields, 0.77 for 3-month yields), but we also see prolonged periods where U.S. interest rates deviate from the weighted average of foreign yields. In particular, since the start of the sample period, 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 GFC, and again prior to the 2020 pandemic recession. All three episodes coincide with Fed monetary tightening cycles. Each period was followed by a U.S. or global recession with U.S. reverting toward foreign yields or vice versa. More 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.

Figure A.2: Foreign Yields and Global Output Growth




The figure plots a correlogram between a foreign yield factor term spread and output growth in the United States, in a sample of advanced economies excluding the U.S., and in a sample of emerging market economies, respectively. The foreign yield factor term spread is constructed as $\mathcal{Y}^{10Y}_{g,t-l} - \mathcal{Y}^{3M}_{g,t-l}$, with the 10-year and 3-month factors as in Equation A.1. Advanced and emerging economies output growth are constructed as GDP-weighted averages of country-specific industrial production series. The dynamic correlations and their 95% confidence bands are computed lagging the foreign factor term spread $l = 0, ...11$ months and correlating against the corresponding industrial production growth measure at $l = 0$. Time 0 reports the contemporaneous correlations. The sample period is January 1999 to December 2018. All variables are monthly frequency.

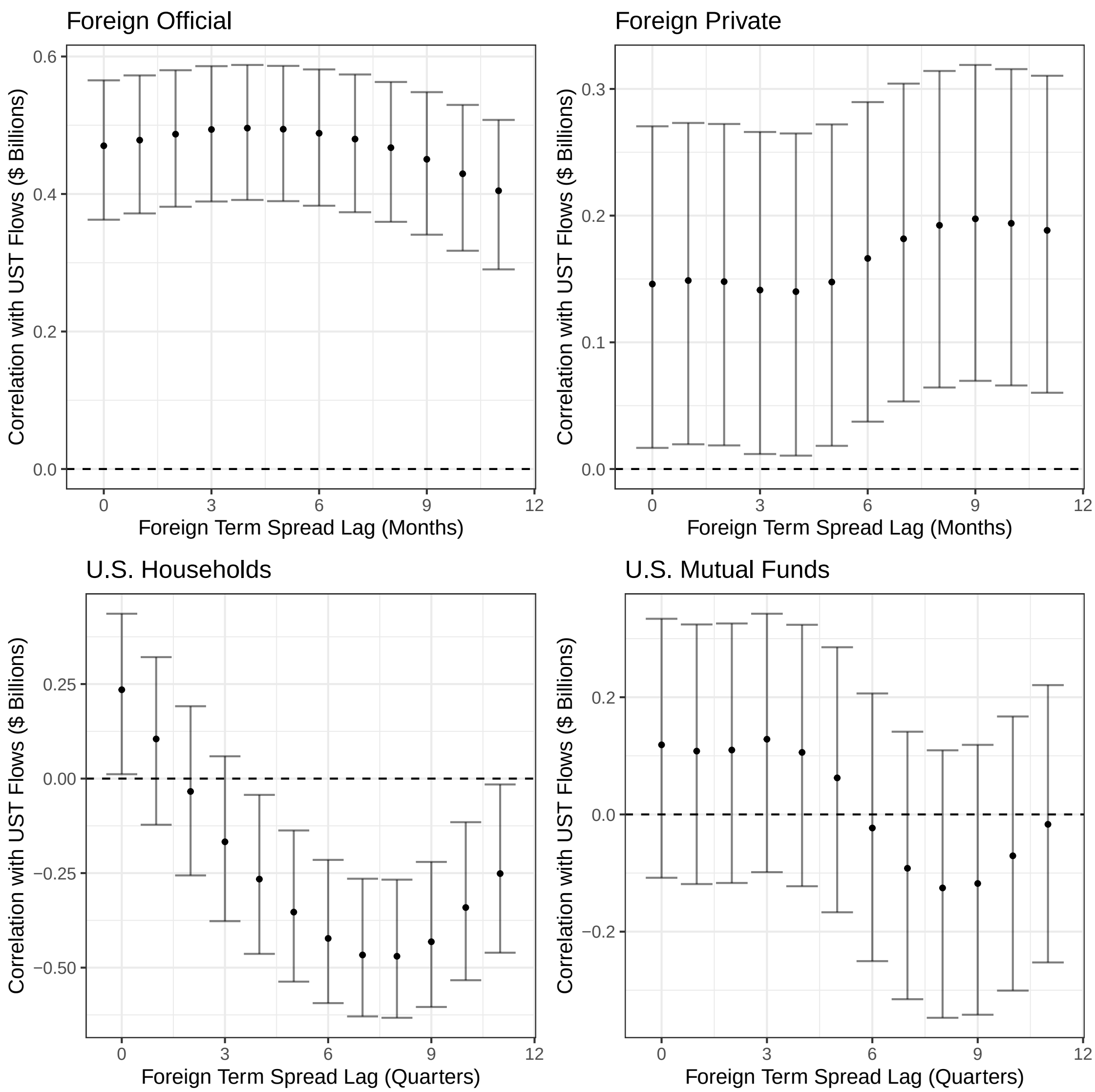
To illustrate and summarize the co-movement of our foreign yield factors, Treasury holdings, and U.S. yields, we now focus on the difference between the 10-year and 3-month common foreign yield factors ($\mathcal{Y}^{10Y}_{g,t} - \mathcal{Y}^{3M}_{g,t}$), which we call the foreign yield factor term spread.[A2] Figure A.2 shows that the foreign yield term-spread factor is significantly and positively correlated with global economic conditions, as measured by industrial production growth in the United States, other advanced economies, and emerging markets. The association becomes much stronger as the time horizon lengthens. In other words, the foreign term-spread factor is associated with current and future output growth.

At the same time, Figure A.3 shows that the foreign term-spread factor is also positively and significantly associated with current and future demand for USTs by foreign *officials* to a much larger degree than foreign *private* UST demand. Recalling that the foreign term-spread factor is strongly procyclical, we also label FO flows pro-cyclical and foreign private flows a-cyclical. In contrast, the foreign term-spread factor is negatively associated with future demand for USTs by domestic U.S. households (including hedge funds), consistent with these investor classes behaving counter-cyclically. Like foreign private demand, U.S. mutual fund demand appears a-cyclical, co-moving weakly with the foreign common term spread.

Figure A.4 provides further detailed analysis on correlations between the foreign term spread and UST demand by sub-groups of countries. Country-level data on UST purchases is the sum of both private and official purchases because public data is unavailable for country-specific official UST purchases. 'Core Economies' refer to countries that make up the foreign yield factors while "Non-Core Economies" are composed of other countries, most of which are Emerging Markets. We also look at correlations based on countries by their exchange rate regime, and we look at purchases from China in isolation. Unsurprisingly, Core Economies have UST purchases that are strongly positively correlated with a steeper foreign term spread, but UST demand from Non-Core Economies is also significantly positively correlated with the foreign term spread. Within Non-Core Economies, fixed exchange rate countries have more pro-cyclical demand for USTs in the short-run, while floating exchange rate countries have a-cyclical demand for USTs in the short-run but pro-cyclical demand for

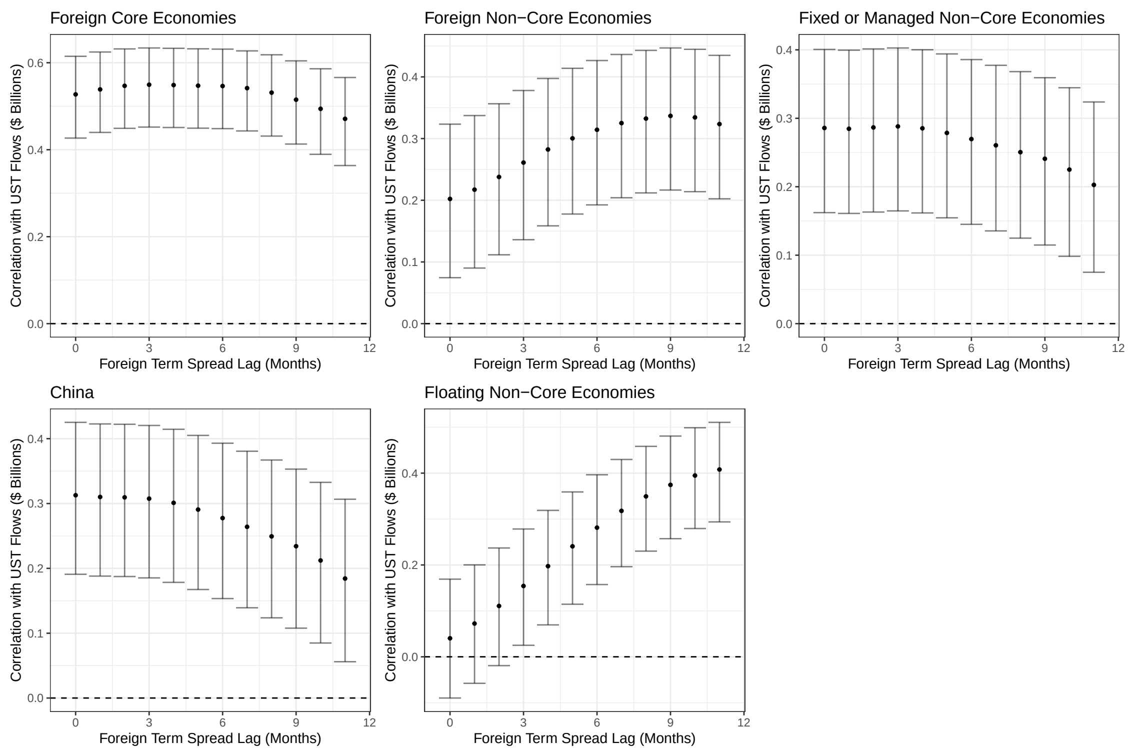
[A2] Looking at common factors in short and long-term yields separately depicts a similar but more complex picture.

Figure A.3: Foreign Yields and UST Net Purchases by Investor Class



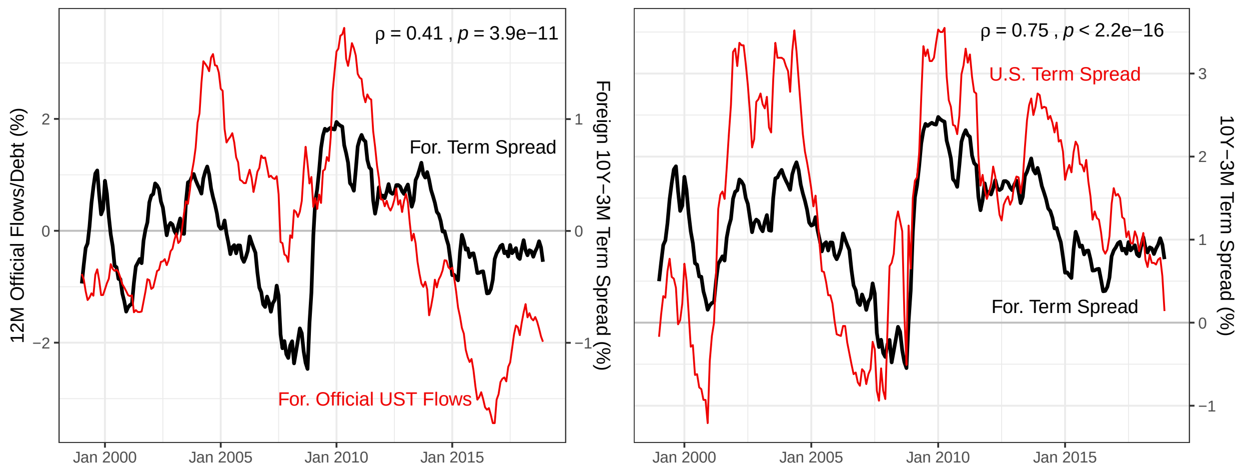

The figure plots a correlogram between a foreign yield factor term spread and 12-month UST flows by the following investor classes: foreign official, foreign private, domestic households (including hedge funds), and domestic mutual funds. The foreign yield factor term spread is constructed as $\mathcal{Y}^{10Y}_{g,t-l} - \mathcal{Y}^{3M}_{g,t-l}$, with the 10Y and 3M factors as in Equation A.1. The dynamic correlations and their 95% confidence bands are computed lagging the foreign factor term spread $l = 0, ...11$ months and correlating against the corresponding UST flows measure at $l = 0$. Time 0 reports the contemporaneous correlations. The sample period is January 1999 to December 2018. Foreign official and private flow data are adjusted benchmark-consistent flows. Foreign UST flows are monthly frequency and U.S. Households and U.S. Mutual Fund UST flows are quarterly.

Figure A.4: Foreign Yields and Foreign UST Net Purchases by Country Type and Exchange Rate Regime




The figure plots a correlogram between a foreign yield factor term spread and 12-month UST flows by the country groups, foreign core (Austria, Australia, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Sweden, Switzerland, United Kingdom), foreign non-core (Brazil, Bulgaria, Chile, China, Colombia, Egypt, India, Indonesia, Israel, S. Korea, Malaysia, Mexico, Panama, Peru, Philippines, Poland, Russia, South Africa, Taiwan, Thailand, Trinidad and Tobago, Turkey, Uruguay, Venezuela, Singapore, Hong Kong), Fixed or managed exchange rate regime against the USD non-core countries (China, Egypt, Panama, Trinidad and Tobago), and floating exchange rate regime non-core countries (Brazil, Chile, Colombia, Israel, S. Korea, Mexico, Poland, South Africa, Thailand, Taiwan, Turkey). Exchange rate regime classifications are based on Ilzetzki et al. [2019]. The foreign yield factor term spread is constructed as $\mathcal{Y}^{10Y}_{g,t-l} - \mathcal{Y}^{3M}_{g,t-l}$, with the 10Y and 3M factors as in Equation A.1. The dynamic correlations and their 95% confidence bands are computed lagging the foreign factor term spread $l = 0, ...11$ months and correlating against the corresponding UST flows measure at $l = 0$. Time 0 reports the contemporaneous correlations. The sample period is January 1999 to December 2018. Foreign UST flow data at the individual country level are monthly frequency, combine official and private flows, and are adjusted benchmark-consistent flows.

Figure A.5: Foreign Yield Spread, FO Flows (left), and U.S. Term Spread (right)




The left panel plots the foreign term spread (thick black line) and the FO flows (thin red line. FO flows is a 12-month rolling sum of adjusted benchmark-consistent data. The foreign term spread constructed as $\mathcal{Y}^{10Y}_{g,t-l} - \mathcal{Y}^{3M}_{g,t-l}$, with the 10Y and 3M factors as in Equation A.1. Both series are de-meaned. Right panel plots the foreign term spread and the U.S. term spread calculated as the 10-year less 3-month U.S. term spread. Both series are not de-meaned. The sample period is January 1999 to December 2018.

USTs in the longer-run. Finally, China exhibits significant, pro-cyclical demand for USTs.

Finally, Figure A.5 illustrates the co-movement between the foreign term-spread factor, FO flows, and the U.S. term spread. Both associations are positive, sizable, and statistically significant. Taken together, suggesting that the foreign term spread may be a common factor jointly correlated with U.S. yields and FO demand for USTs.

These co-movements are consistent with optimistic global growth pushing up interest rates in the U.S. and internationally (Figure A.2), and, at the same time, increasing foreign official demand for USTs (Figure A.3). Higher foreign official demand may stem from stronger export demand during global expansions and/or higher capital inflows prompting foreign officials to accumulate international reserves, especially in countries where monetary policy limits exchange rate flexibility. Domestic private investors, in contrast, seem to rotate out of USTs as risk appetite increases during booms (Figure A.3). Conversely, during economic downturns or episodes of global financial distress, interest rates fall in the U.S. and other advanced economies (Figure A.5), as U.S. domestic private investors may seek safety, while foreign officials, mostly emerging markets, tend to liquidate international reserves to stabilize exchange rates. Indeed, Dominguez et al. [2012] documents 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.

Taken together, therefore, the documented co-movements are consistent with the hypothesis that foreign officials reserve managers accumulate USTs pro-cyclically as rates rise internationally, while domestic private investors behave counter-cyclically buying USTs when global rates are falling and foreign officials are selling. As a consequence, omitting foreign common yield factors, which positively co-vary with foreign official UST demand and U.S. yields, can bias estimates of the effect of FO flows on U.S. yields as we illustrate in the paper both in theory and in the data.

## A3 VAR Residual Heteroskedasticity and Non-normality

This section presents additional evidence on the reduced form VAR residuals from the estimation of VAR in (6) consistent with the assumptions made for identification via heteroskedasticiy.

Table A.2: Higher Moments of the Reduced-Form VAR Residuals

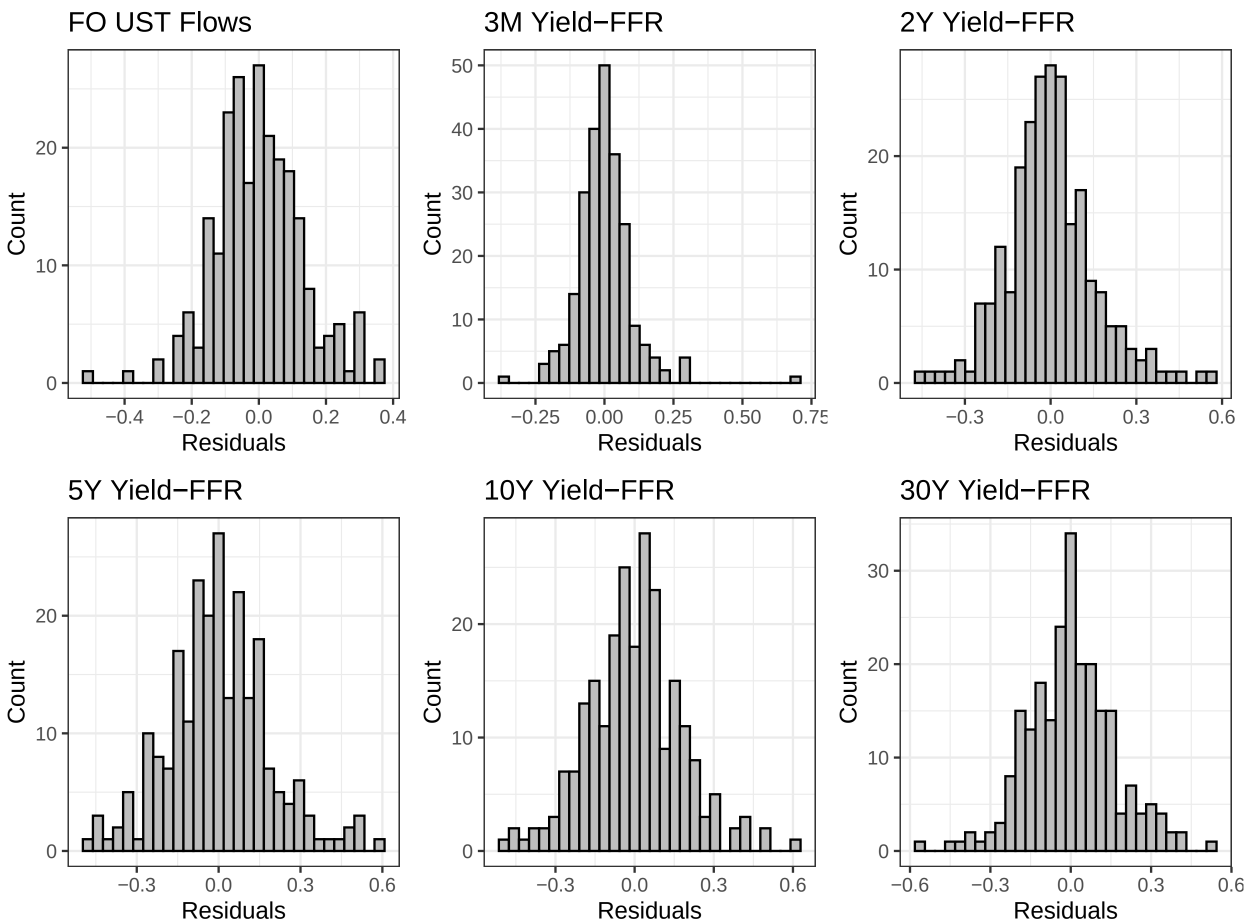
| | $\Delta FO_t$ | $y_{us,t}^{3M-FF}$ | $y_{us,t}^{2Y-FF}$ | $y_{us,t}^{5Y-FF}$ | $y_{us,t}^{10Y-FF}$ | $y_{us,t}^{30Y-FF}$ |
|---|---|---|---|---|---|---|
| 1st-Order Autocorrelation | 0.01 | 0.15 | 0.03 | 0.05 | 0.06 | 0.04 |
| SW Normality Test (p-value) | 0.01 | 0.00 | 0.00 | 0.04 | 0.12 | 0.07 |
| 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 |

The table report higher moment summary statistics for residual series of the reduced form VAR in Equation 6. The SW Normality test refers to the Shapiro-Wilk test, where the null hypothesis is that the variable is normally distributed.

Table A.2 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.6. A critical assumption for our identification strategy through heteroskasticity is the presence of time-varying volatility in the residuals. 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 with the rejection of the variance ratio tests in Section 4.

Figure A.6: Histograms of Reduced-Form VAR Residuals




The figure plots histograms of the residual series from reduced form VAR in Equation 6.