Convenience yields and the foreign demand for US Treasuries

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Abstract

This paper investigates the role of convenience yields in determining the yield sensitivity of the foreign demand for US Treasuries and the equilibrium interest rates of government bonds. I build a portfolio choice model featuring an investor with standard mean-variance prefernces (banks), and another that derives a non-monetary payoff from holding US Treasuries (insurances). The model can explain why insurances hold US Treasuries even if they offer negative excess returns and poor hedging properties against income risk. It predicts that preferences for non-monetary payoffs reduce yield sensitivity for a given level of risk aversion, and that equilibrium excess returns react less strongly to debt supply in the presence of convenience yield investors. Structural parameters recovered from an estimation of the model on data from European banks and insurances reveal that the preference for non-monetary payoffs reduces the yield sensitivity of insurances by 60%. It also explains 50% of the difference in the sensitivity to excess returns across sectors. However, it accounts for only 2% of the decline in Treasury excess returns in response to an increase in eurozone debt supply.

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1 Introduction

US Treasuries are the premier global safe asset, and their special role affords them a premium, or convenience yield, reflected in lower returns compared to other advanced economy sovereign debt (Du et al., 2018) and other dollar-denominated assets with similar safety and liquidity features (Longstaff, 2004, Krishnamurthy and Vissing-Jorgensen, 2012).

The convenience yield of US Treasuries is crucial for the sustainability of the burgeoning US government debt, as it allows the US government to borrow more cheaply than sovereigns with comparable credit rating, and it can explain the gap between the market value of US debt and projected government deficits (Jiang et al., 2019). This funding advantage is driven in large part by the inelastic demand for Treasuries by foreign investors, who are willing to accept a lower yield to meet their need for safe and liquid dollar-denominated assets (Jiang et al., 2022).

Among the key foreign investors contributing to convenience yields, two groups stand out: the official sector, which holds Treasuries as a crucial part of foreign reserves(Alfaro et al., 2014, Ilzetzki et al., 2017), and, of particular relevance to this study, insurance and pension funds (ICPF). Due to the nature of the insurance business, this sector needs safe and liquid assets with stable payoffs, especially at the long end of the yield curve due to duration matching between assets and liabilities (Greenwood and Vayanos, 2010, Greenwood and Vissing-Jorgensen, 2018). Japanese ICPFs play a particularly central role as channelers of large amounts of savings into foreign safe and liquid assets in the face of limited domestic investment opportunities. This mechanism came to the fore especially in the context of the "saving glut" in Asian economies in the early 2000's (Bernanke, 2005, Bernanke et al., 2010, Barsky and Easton, 2021).

At the same time, there is substantial heterogeneity in the yield elasticity of safe asset demand across investor categories. Tabova and Warnock (2022) singles out the foreign official sector as particularly inelastic in its demand for Treasuries, while foreign private investors are sensitive to yields. Within private investors, Fang et al. (2023) finds that non-banks, including insurance and pension funds (ICPF), absorb a large amount of sovereign debt issuance, and display a particularly low yield elasticity for advanced-economy debt. On the contrary, the banking sector is more responsive to yields than insurances, for both assets in general (Timmer, 2018), and US sovereign debt specifically (Eren et al., 2023). [1] Consistently with my findings, Koijen et al. (2021) report that, in the euro area, the yield elas-

[1] Note that (Eren et al., 2023) break down the ICPF sector into pension funds and insurances, finding a slightly larger yield elasticity than commercial banks for the former, but a much lower

ticity of banks' demand for European government bonds is amongst the highest across sectors, while European insurance companies even have a *negative* elasticity.

The literature explains the differences in demand elasticity mostly in terms of risk management practices (Eren et al., 2023), regulatory framework (Faia et al., 2022), or market-making versus speculative roles (Abbassi et al., 2016, Timmer, 2018). Instead, this paper provides a theoretical framework that motivates the heterogeneity in yield sensitivity through the lens of preferences for non-monetary payoffs. I zoom in on the difference between insurances and banks, as the former have been identified as an important driver of the convenience yield; and the latter are among the most yield-sensitive sectors.

The theoretical framework consists of a simple mean-variance model of portfolio choice between US and domestic government bonds, in which the insurance sector derives a non-monetary payoff from US Treasuries, while banks have standards preferences. It shows that the heterogeneity in sensitivity across sectors can be due either to differences in risk aversion, or to the preference for non-monetary payoffs. I model preference for the special safety and liquidity of US Treasury as an additional term in the investor's objective function, following the approach of Krishnamurthy and Vissing-Jorgensen (2012) and Nagel (2016), among others.

The model predicts that insurances have a lower sensitivity to excess returns on Treasuries than they otherwise would in the absence of convenience yields. Furthermore, the optmial portfolio share for insurances implies that this sector would be willing to hold a non-zero amount of US Treasuries even if they both pay a lower return than domestic bonds, and do not provide a good hedge for income risk. This feature emerges uniquely from the convenience yield mechanism and cannot be explained by risk aversion: investors that value solely monetary payoffs hold assets only if they deliver an excess returns, or if they are a good hedge.

Equilibrium excess returns depend on the relative supply of US and domestic government debt, and they can be decomposed in a risk premium term and a convenience yield term. The presence of "convenience yield" investors drives down excess returns and makes them more sensitive to changes in debt supply, in line with the safe asset supply channel of quantitative easing highlighted in Jiang et al. (2021b) and Christensen et al. (2023).

The portfolio equations for the two sectors, jointly with the expression for equilibrium returns, imply that exogenous changes in debt supply are a valid instru-

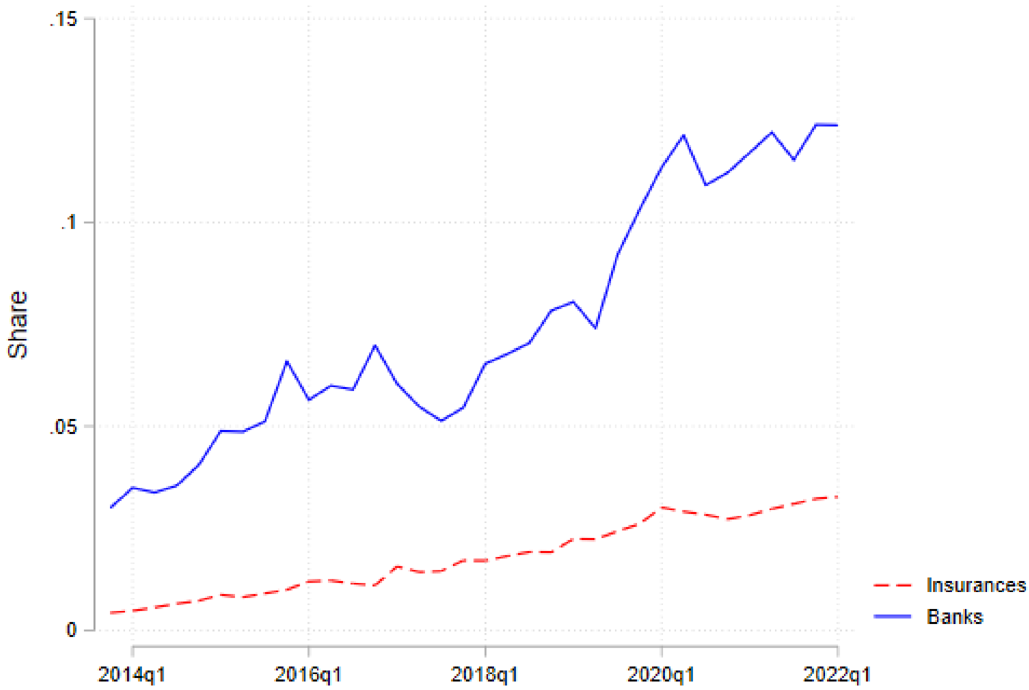
and not statistically significant response for the latter.

ment for excess returns in the portfolio equations using a two-stage least squares (2SLS) procedure. The restrictions on the estimated first- and second-stage coefficients then allow to back out the structural parameters regulating risk aversion and the preference for non-monetary payoffs. The structural parameters are crucial in disentangling and quantifying the role of risk aversion and convenience yields, as the predictions on demand sensitivity concern counterfactuals, rather than a direct comparison of observables across sectors.

The underlying rationale for the different demand sensitivity is crucial for the capacity of markets to absorb additional US government debt. If the yield-insensitive demand by foreigners is due to risk aversion, the elasticity is heavily dependent on contingent market developments as encapsulated by the variance of US returns. Therefore, events such as a temporary uncertainty on fiscal sustainability could jeopardise the ability of the US government to fund its debt cheapy. Conversely, convenience yields are tightly linked to the status of the US dollar as reserve currency, and of US Treasuries as global safe assets. These are much more persistent phenomena (Coppola et al., 2023), liable to evolve only in the face of extreme events such as a default (Choi et al., 2024) or major geopolitical upheaval (Eichengreen and Flandreau, 2009). Therefore, it is lkely that low yield sensitivity is the more stable, and the more reliable from the point of view of the US government, the more it is driven by preference for the special features of Treasuries.

I estimate the model on data from the banking and insurance sector in the eurozone. The reason for this geographical focus is twofold. Firstly, the Public Sector Purchase Programme (PSPP) by the European Central Bank (ECB) generates exogenous variation in the supply of sovereign debt issued by euro zone countries, solely as a function of the ECB's Capital Key and of the maturity structure of outstanding government bonds under the principle of market neutrality. This feature, combined with cross-country difference in convenience yields of euro-denominated government bonds documented in Jiang et al. (2020), provide an ideal laboratory to study responses to excess returns driven by exogenous changes in the relative supply of safe assets with an instrumental variables approach that matches the sets of equations derived from the theoretical model. This identification strategy is also exploited by Koijen et al. (2021) in the setting of demand for European government bonds. Secondly, the very similar regulatory framework for banks and insurances in the realm of sovereign bonds and exposure to foreign exchange risk removes a potential alternative explanation for cross-sector differences in demand elasticity, thus sharpening the focus on differences in preferences. The approach of estimating a mean-variance portfolio model through instrumental variables is consistent with the framework laid out in Koijen and Yogo (2019). However, among other differences, I do not specify the full demand system, and I solve for equilibrium returns as a function of debt supply rather than relying on a factor model. Furthermore, this paper makes a step in the direction of understanding the nature of demand heterogeintiy at the core of the Koijen and Yogo (2019) model.

Figure 1. US sovereign portfolio share for banks and insurances



Share of US Treasury holdings in a portfolio including US Treasuries and euro area government bonds for all banks (solid blue lines) and insurance and pension funds (dashed red line) domiciled in the euro area Source: European Central Bank Securities Holdings Statistics database.

Figure 1 plots the share of US Treasuries in a portfolio including euro area government bonds for banks (solid blue line) and ICPF (dashed red line) domiciled in the euro area. It is evident at first glance that the banks' share is much more volatile. While the comparison of unconditional volatility is not enough to draw any conclusions, it is certainly suggestive as to the plausible lower sensitivity of insurances' Treasury holdings to excess returns.

Figure 2 shows the correlation between the excess returns of Treasuries with respect to bonds issued by a given eurozone country, and the income of banks and insurance companies resident in the same country. The correlation is negative for banks and positive for insurances. Therefore, by this measure US government bonds are not a good hedge for the income risk of insurers. In the period between 2011 Q4 and 2023 Q3 over which the graph is constructed, Treasuries offer on average *negative* excess returns of about a quarter of a percentage point. [2] [3] Therefore, insurance companies would have no incentive to include Treasuries in their portfolio under standard preferences that value assets solely for the balance of risks and rewards in monetary returns. This apparent contradiction can be resolved by the model presented in this paper: the presence of non-monetary payoffs motivates insurances to hold US Treasuries even in the face of a poor risk-return trade-off. Therefore, the observation in Acharya and Laarits (2023) that US Treasuries earn convenience yields because of their hedging properties againt stock market risk does not appear to extend to the case of income risk for insurances.

Estimates of the portfolio equations via 2SLS reveal that banks increase their US Treasury portfolio share by 6.93 percentage points in response to a one percentage point increase in the convenience yield component of excess returns brought about by exogenous changes in the supply of eurozone governmnet debt. In contrast, insurances increase their portfolio share by only 2.31 percentage points. These findings are in line with existing evidence of lower sensitivity to excess returns of insurances' demand for government debt.

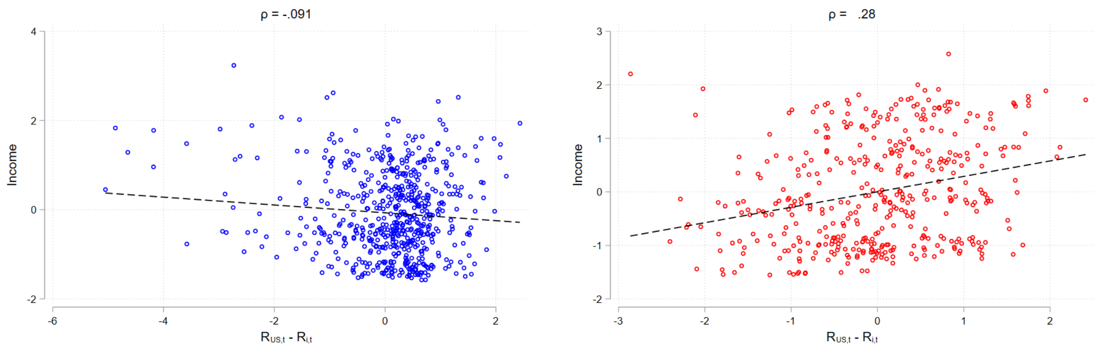
The structural parameters recovered from this estimation procedure imply that banks are about ten times more risk tolerant than insurances, and that non-monetary payoffs of Treasuries have a 10% weight in insurances' preferences. To understand the implications of these parameters for both portfolio choice and equilibrium interest rates, I perform three counterfactual experiments in the model.

First, I show that absent convenience yields insurances would display a 60% higher sensitivity to excess returns. Then, I calculate that 50% of the difference in sensitivity between banks and insurers is attributable to preference for the non-

[2] Table A2 in appendix B reports summary statistics for the variables used in the empirical analysis

[3] Here it is important to clarify the notion of excess returns used. In the model, US Treasury excess returns arise solely due to exogenous, stochastic fluctuations in exchange rates. Therefore, I account for the exchange rate expectations term in excess returns by adjusting for forward rates. This proxy is nonetheless imperfect due to the documented deviations in Covered Interest Parity stemming from frictions in FX markets (Borio et al., 2016, Rime et al., 2022). Other sources of variation that are disregarded in the model but likely affect the data, like sovereign credit risk, are allowed to influence this measure.

Figure 2. Correlation between Treasury excess returns and income



(a) Banks (b) Insurances

Correlation between excess returns on US Treasuries compared to country i's government bonds, and the income of banks (left-hand side panel) or insurance companies (right-hand side panel) domiciled in country i. The correlation is calcu lated over the country-quarter distribution on data from 2011 Q4 to 2023 Q2 for all euro area countries excluding Greece. Excess returns are averaged over the 1,2,3,5, and 10 year maturities and over quarters, and are adjusted for exchange rate forward premia as a market-implied measure of expected changes in the exchange rate of the euro *vis à vis* the dollar. Income is calculated as total income for banks, and total income from premia for insurances. Sources: Refinitiv Eikon, European Central Bank Consolidated Banking database, and EIOPA Insurance Statistics.

monetary payoffs of Treasuries. Finally, I decompose the response of equilibrium excess returns to changes in debt supply into a risk premium and a convenience yield component, and quantify the contribution of the latter as only 2%.

Therefore, the structural parameters recovered by combining the theoretical model and the data reveal that the convenience yield is a quantitatively important determinant of the portfolio choice for a class of investors, and it can explain a large portion of the observed cross-section difference in the yield sensitivity of sovereign portfolio shares. On the other hand, the reaction of equilibrium interest rates to changes in debt supply does not seem to be substantially affected by the presence of investors that value non-monetary payoffs. Although limited to the context of banks and insurances in Europe, this result raises questions on the potency of the safe asset channel of quantiative easing, and on the wisdom of relying on convenience yields to fund government debt cheaply.

The remainder of the paper is organised as follows. Section 2 lays out a simple mean-variance portfolio problem with convenience yields, and derives propositions on portfolio shares and equilibrium excess returns. Section 3 estimates the coefficients of a linearised version of the model via 2SLS on data from eurozone banks and insurances. Section 4 recovers the structural parameters on investors' preferences from the estimated coefficients, and runs counterfactual experiments within the model. Section 5 concludes.

2 A model of portfolio choice with convenience yields

The evidence presented in the introduction suggests that the motives behind foreign investors' holdings of US Treasuries might vary across sectors. In this section, I build a simple portfolio choice model to explain the difference in behaviour in terms of heterogeneous preferences over US Treasuries, and derive the implications for asset pricing and the yield sensitivity of demand.

I model the static choice between euro-denominated government bonds, offering a deterministic return, and US Treasuries, whose payoff is stochastic due to exogenous exchange rate fluctuations that are not modelled directly. Therefore, I implicitly assume no hedging of exchange rate risk. Importantly, the model also abstracts from sovereign credit risk, as the focus is on differences in returns that are motivated by convenience yields. Since credit risk is non-negligible for several countries in my sample, I control for it in the empirical analysis via credit default swaps (CDS) rates and country fixed effects. Investors also receive a stochastic income, which represents revenues from all other activities, for example loans for banks and premia from insurance. This assumption aims to capture succinctly other sources of income that are outside of the scope of this model of government bond portfolios, but nonetheless affect the investment choice by virtue of their correlation with sovereign returns.

Two agents populate the model: "conventional" investors have standard Markowitz (1952) mean-variance preferences over wealth, while "convenience yield" investors additionally derive utility from holding US government bonds. This approach for modelling convenience yields is standard in the literature, and the utility is justified with the special liquidity and/or safety features of US Treasuries as the global safe asset .[4] In keeping with the empirical section of the paper, I conceptualise the two classes of investors as banks and insurances, respectively.

[4] A non-exhaustive list of papers that adopt this approach includes Krishnamurthy and Vissing-Jorgensen (2012), Engel (2016), Engel and Wu (2018), **?**, Jiang et al. (2021b), Nagel (2016), Jiang et al. (2021a), and Bodenstein et al. (2023).

2.1 Banks

Banks choose to allocate their initial wealth W_0^B between euro area government bonds b^B and US Treasuries b_{US}^B. The banks' problem is

$$\begin{aligned} &\max_{b^B, b_{US}^B} \mathbb{E}[W^B] - 0.5\gamma^B \mathbb{V}[W^B] \\ \text{s.t.} \quad & W^B = R_b^B + R_{US} b_{US}^B + Y^B \\ & b^B + b_{US}^B = W_0^B, \end{aligned}$$

where $\gamma^B > 0$ is the banks' risk aversion parameter, W^B is their final wealth, Y^B is their stochastic income, R is the deterministic return from euro area government bonds, and R_{US} is the stochastic return from US Treasuries.

The first-order condition for US Treasury holdings is

$$b_{US}^B = \frac{1}{\gamma^B \mathbb{V}[R_{US} - R]} \left(\mathbb{E}[R_{US} - R] - \gamma^B Cov[R_{US} - R, Y^B] \right), \tag{1}$$

which is the standard formula for asset allocations in a Markowitz (1952) model. This condition implies that either $\mathbb{E}[R_{US} - R] > 0$ or $\gamma^B Cov[R_{US}, Y^B] < 0$ (or both) are required for $b_{US}^B > 0$. Therefore, "conventional" investors choose to hold a positive amount of US Treasuries only if they offer either a positive excess return, or insurance against income risk. Contrary to "convenience yield" investors, in this model banks only value US Treasuries for their risk-return profile, just as they would any other asset.

To obtain an expression for the portfolio share $s_{US}^B := \frac{b_{US}^B}{b_{US}^B + b_{US}}$, note that $b^B + b_{US}^B = W_0^B$ implies $s_{US}^B = \frac{1}{W_0^B \gamma^B \mathbb{V}[R_{US} - R]} \left(\mathbb{E}[R_{US} - R] - \gamma^B Cov[R_{US}, Y^B] \right)$. Thus, letting $W_0^B = 1$ without loss of generality, it follows that $s_{US}^B = b_{US}^B$.

The sensitivity of banks' US Treasury portfolio share to excess returns is

$$\frac{\partial s_{US}^B}{\partial \mathbb{E}[R_{US} - R]} = \frac{1}{\gamma^B \mathbb{V}[R_{US} - R]}. \tag{2}$$

It is immediate to see that $\frac{\partial s_{US}^B}{\partial \mathbb{E}[R_{US} - R]} > 0$ for $\gamma^B > 0$. Therefore, standard mean-variance preferences predicts a low sensitivity to excess returns, through high risk-aversion or high variance of returns. In the next section, I demonstrate how this statement changes in the case of investors that derive a non-monetary payoff from US Treasuries.

2.2 Insurances

Insurances choose to allocate their initial wealth W_0^I between euro area government bonds b^I and US Treasuries b_{US}^I. Unlike banks, insurances derive utility directly from holding US Treasuries, formalised as an additive term to their otherwise standard mean-variance objective function. The insurances' problem is

$$\begin{aligned} \max_{b^I, b_{US}^I} \ & \mathbb{E}[W^I] - 0.5\gamma^I \mathbb{V}[W^I] + \psi \log(b_{US}^I) \\ \text{s.t.} \quad & W^I = R_b^I + R_{US} b_{US}^I + Y^I \\ & b^I + b_{US}^I = W_0^I, \end{aligned}$$

where $\gamma^I > 0$ is the insurances' risk aversion parameter, W^I is their final wealth, Y^I is their stochastic income, and ψ is a parameter regulating the weight of non-monetary payoffs in the objective function. The log functional form for the non-monetary payoff term facilitates calculations, but the predictions of the model would not change, qualitatively, by using other increasing, concave funcitons instead.

By taking the first-order condition for b_{US}^I, and letting $W_0^I = 1$ as in the banks' problem, I can derive the following proposition on the optimal US Treasury portfolio share for insurers:

Proposition 1 (Optimal portfolio share for insurances). (i) *The optimal portfolio share is*

$$s_{US}^I = \frac{\mathbb{E}[R_{US} - R] - \gamma^I Cov[R_{US} - R, Y^I]}{2\gamma^I \mathbb{V}[R_{US} - R]} + \frac{\sqrt{(-\mathbb{E}[R_{US} - R] + \gamma^I Cov[R_{US} - R, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US} - R]}}{2\gamma^I \mathbb{V}[R_{US} - R]}. \tag{3}$$

(ii) $s_{US}^I \in \mathbb{R}^+$ *if* $\gamma^I > 0$, $\psi > 0$, *and* $\mathbb{V}[R_{US} - R] > 0$.

Proof. In Appendix C.1. □

Proposition 1 (i) shows that the optimal portfolio share for insurers can be decomposed in two parts that have a natural interpretation: the first is isomporphic to the standard mean-variance optimal share, and reflects the risk-return tradeoff of Treasuries; the second depends on ψ and represents the additional motive to hold Treasuries for investors that derive a convenience yield from them.

Note that, for $\gamma^I > 0$, $\psi > 0$, and $\mathbb{V}[R_{US} - R] > 0$, $s^I_{US} > s^I_{US}|_{\psi=0}$. Due to their preference for non-monetary payoffs, insurances have a further motive to hold Treasuries in addition to their risk-return profile. Therefore, in equilibrium they choose a higher portfolio share than they would absent convenience yields.

Furthermore,

$$\lim_{\psi \to 0} s^I_{US} = \frac{1}{\gamma^I \mathbb{V}[R_{US} - R]} \left(\mathbb{E}[R_{US} - R] - \gamma^I Cov[R_{US} - R, Y^I] \right).$$

Therefore, the optimal share collapses back to the standard case as the weight on non-monetary payoffs vanishes.

Proposition 1 (ii) states the conditions under which the insurers' problem admits a real, positive solution for s^I_{US}. Note that there are no requirements on the risk-return profile of US Treasuries. Therefore, $b^I_{US} > 0$ even for $\mathbb{E}[R_{US} - R] < 0$ and $Cov[R_{US} - R, Y^I] > 0$ simultaneously. Due to the non-monetary payoff of Treasuries, insurers choose to hold a positive amount even if they offer neither an extra return, nor good insurance for income risk. The model can thus rationalise the puzzling joint observation of positive US Treasury portfolio shares for insurers; negative Treasury excess returns; and a positive correlation between insurers' income and Treasury returns.

To understand how the preference for Treasuries affects the yield sensitivity of portfolio shares, consider the following proposition:

Proposition 2 (Sensitivity of insurances to excess returns)**.** (i)

$$\frac{\partial s^I_{US}}{\partial \mathbb{E}[R_{US} - R]} = \frac{1}{2} \frac{1}{\gamma^I \mathbb{V}[R_{US} - R]} \left(1 + \frac{-\mathbb{E}[R_{US} - R] + \gamma^I Cov[R_{US} - R, Y^I]}{\sqrt{(-\mathbb{E}[R_{US} - R] + \gamma^I Cov[R_{US} - R, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US} - R]}} \right) \tag{4}$$

(ii) $\frac{\partial s^I_{US}}{\partial \mathbb{E}[R_{US} - R]} \in \left(0, \frac{1}{\gamma^I \mathbb{V}[R_{US} - R]}\right)$ *for* $-\mathbb{E}[R_{US} - R] + \gamma^I Cov[R_{US}, Y^I] > 0$*,* $\gamma^I > 0$*,* $\psi > 0$*, and* $\mathbb{V}[R_{US} - R] > 0$*.*

Proof. In Appendix C.2. □

Therefore, the presence of convenience yields results in a *lower* sensitivity of insurers' US Treasury portfolio shares to excess returns compared to a situation in which there are no non-monetary payoffs, as shown by proposition 2 (ii). However, the predicted reaction to excess returns is still always positive. Note that the condition $-\mathbb{E}[R_{US} - R] + \gamma^I Cov[R_{US}, Y^I] > 0$ that ensures $\frac{\partial s^I_{US}}{\partial \mathbb{E}[R_{US}-R]} > 0$ is compatible with the empirically relevant case $\mathbb{E}[R_{US} - R] < 0$ and $Cov[R_{US}, Y^I] > 0$, which implies that US Treasuries offer neither a positive excess return nor a good hedge for insurers' income risk.

A higher risk aversion parameter γ^I would also result in a lower, but positive, sensitivity to excess returns. In order to quantify the relative importance of the two effects, it is then crucial to estimate the structural parameters γ^I and ψ.

To understand the role of non-monetary payoffs in the sensitivity to excess returns, it is instructive to inspect the limits of $\frac{\partial s^I_{US}}{\partial \mathbb{E}[R_{US}-R]}$ for ψ.

$$\lim_{\psi \to 0} \frac{\partial s^I_{US}}{\partial \mathbb{E}[R_{US} - R]} = \frac{1}{\gamma^I \mathbb{V}[R_{US} - R]}$$

Therefore, as the weight of non-monetary payoff in the insurer's preferences vanishes, the sensitivity of the US Treasury portfolio share to excess returns approaches the standard mean-variance case.

$$\lim_{\psi \to \infty} \frac{\partial s^I_{US}}{\partial \mathbb{E}[R_{US} - R]} = 0$$

On the other hand, as the weight goes to infinity the sensitivity to excess returns tends to zero: the more important non-monetary payoffs are in the investors' preferences, the less they react to excess returns, as the latter become a relatively irrelevant feature of the asset.

2.3 Equilibrium and pricing

The previous sections analysed the response of banks' and insurances' US Treasury portfolio share to changes in excess returns, while remaining agnostic on the source of the latter. In this section, I derive excess returns in equilibrium as a function of the relative supply of euro area and US bonds. Thus, I obtain a theoretical counterpart for the empirical identificaiton strategy, which exploits exogenous changes in the supply of euro area government securities.

Equilibrium excess returns

The market clearing conditions for euro area and US government bonds, respectively, are

$$
\begin{aligned}
b^I + b^B &= B \\
b^I_{US} + b^B_{US} &= B_{US},
\end{aligned}
$$

Where B is the supply of euro area bonds, and likewise B_{US} is the supply of US Treasuries. I then define an equilibrium as a set of portfolio allocations $\{b^I, b^B, b^I_{US}, b^B_{US}\}$ and Treasury excess returns $\mathbb{E}[R_{US} - R]$ for which the first-order conditions of banks and insurances hold, and the markets for euro area and US government bonds clear.

To derive equilibrium expected excess returns, sum the first-order conditions of both investors, defining for ease of exposition $\tau^X := \frac{1}{\gamma^X}$, the risk tolerance parameter of investor $X = \{B, I\}$.

$$
\begin{aligned}
\mathbb{E}[R_{US} - R] = &\underbrace{\frac{\mathbb{V}[R_{US} - R]}{\tau^B + \tau^I} B_{US} + \frac{Cov[R_{US} - R, Y^B] + Cov[R_{US} - R, Y^I]}{\tau^B + \tau^I}}_{\text{Risk premium} := RP} \\
&\underbrace{- \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{b^I_{US}}}_{\text{Convenience yield} := \phi}
\end{aligned}
\tag{5}
$$

The equilibrium excess return of US Treasuries comprises two parts that can be interpreted intuitively. The first is a standard risk premium term: increasing in the volatility of excess returns and in the covraiances between excess returns and income; and decreasing in the investors' risk tolerance. The second is specific to this model, and it can be interpreted as a convenience yield. The higher the weight ψ of non-monetary payoffs in insurers' preferences, the lower the equilibrium excess returns *ceteris paribus*. Since Treasuries offer non-monetary payoffs to certain investors, they are willing to accept a lower monetary return, and this is is reflected in equilibrium. This mechanism is well-studied in the literature on US Treasury pricing, since at least Krishnamurthy and Vissing-Jorgensen (2012), and it can explain the observed premium on US Treasuries (Du et al., 2018). Note that the importance of the convenience yield term depends on the relative risk tolerance of insurances and banks. A higher risk tolerance of insurances implies a larger weight of their preferences on the equilibrium excess returns, and so a larger convenience yield term.

Much like the deviations from interest parity arising in open-economy macroeconomic models that incorporate convenience yields, the presence of Treasury holdings in the payoff function introduces a wedge in the pricing equation (Engel and Wu, 2018, Jiang et al., 2021a, Valchev, 2021). However, in this model Treasuries carry exchange rate risk from the perspective of European investors, so the usual interest parity condition does not generally hold even in the absence of "convenience yield" investors. As a consequence, the observed negative excess returns of Treasuries could be explained by a strongly negative correlation between excess returns combined with a relatively low excess return volatility. However, in the data we observe $Cov[R_{US} - R, Y^I] > 0$ with a similar magnitude as the negative $Cov[R_{US} - R, Y^B]$. Therefore, from the perspective of the asset pricing equation implied by the model, the risk-return profile of US Treasuries for European investors is not likely to be a convincing explanation for negative excess returns.

The effects of debt supply

Equilibrium excess returns depend on debt supply through the risk premium component: the higher the amount of risky asset B_{US}, the higher the risk premium required for investors to absorb it. However, quantities enter equation 5 through the convenience yield term too, specifically in the form of insurers' Treasury holdings B^I_{US}. Due to the diminishing marginal utility of US government bonds, the weight of the convenience yield term is decreasing in the amount held by insurances. This result is discussed in Jiang et al. (2021b), which shows in a general equilibrium model how central bank quantitative easing affects asset prices also by altering the relative supply of safe assets with non-monetary payoffs.

Re-write Equation 5 as a function of the supply of euro area governemnt bonds B, using the market clearing conditions and the budget constraints of both agents.

$$\begin{aligned}\mathbb{E}[R_{US} - R] =& \frac{\mathbb{V}[R_{US} - R]}{\tau^B + \tau^I}(W_0^I + W_0^B - B) + \frac{Cov[R_{US} - R, Y^B] + Cov[R_{US} - R, Y^I]}{\tau^B + \tau^I} \\ & - \frac{\tau^I}{\tau^B + \tau^I}\frac{\psi}{W_0^I - B + b^B}.\end{aligned}$$

Excess returns then depend on the supply of euro area government bond B through both the risk-premium and the convenience yield terms. A change in B alters not only the relative amount of safe versus risky assets on the market, hence affecting the risk premium, but also the relative amount of US Treasuries that insurers absorb, hence affecting the equilibrium convenience yield.

In the empirical estimation, I exploit exogenous changes in the supply of European government bonds due to the implementation structure of the PSPP programme by the European Central Bank. I will use PSPP holdings as an instrument for changes in the convenience yield portion of excess returns in a two-stage least squares setup. To undestand the model-implied sign of the coefficient on PSPP amounts in the first-stage regression, I analyse the derivative of $\mathbb{E}[R_{US} - R]$ with respect to B. Furthermore, this derivative implies restrictions on ψ, τ^B and τ^I that I exploit to back out these structural parameters. Excess returns are the simple sum of the risk premium RP and the convenience yield ϕ. So, I compute the derivatives of the two components separately, and the derivative of the excess return is the sum of the two. The derivative of the convenience yield term is of particular importance, as it is the object of the empirical analysis.

Proposition 3 (Reaction of excess returns to euro area debt supply). (i)

$$\frac{\partial RP}{\partial B} = \frac{-\mathbb{V}[R_{US} - R]}{\tau^B + \tau^I} < 0. \tag{6}$$

(ii)

$$\frac{\partial \phi}{\partial B} = \frac{\tau^I \mathbb{V}[R_{US} - R]}{(\tau^B + \tau^I)\mathbb{V}[R_{US} - R](W_0^I - B + b^B)^2 + \tau^B \tau^I \psi} \left(\frac{\tau^B}{\tau^B + \tau^I} - 1 \right) < 0. \tag{7}$$

Proof. In Appendix C.3. □

Proposition 3 (ii) then predicts a negative reaction of US Treasuries excess returns through the convenience yield component. An increase in the supply of euro area government bonds reduces the relative amount of US Treasuries that investors have to absorb in equilibrium. Due the concavity of non-monetary payoffs, insurers are now willing to accept a lower return *ceteris paribus* on Treasuries as their relative supply decreased. [5] The empirical model uses PSPP purchases, which correspond to a *decrease* in the market supply of euro area governemnt bonds. Therefore, proposition 3 (ii) predicts a *positive* coefficient in the first-stage regression of the convenience yield component of excess returns on PSPP holdings.

[5] I implicitly assume that changes in B have no effect on variances and covariances, which are treated as fixed parameters.

3 Estimation

The model makes three predictions on the sensitivity of portfolio shares to excess returns, and on the reaction of excess returns to changes in debt supply. First, banks react to an increase in the excess return of US over euro area government bonds by increasing their US Treasury sovereign portfolio share. Second, insurances increase their US Treasury sovereign portfolio shares in response to higher excess returns, but by a smaller amount than they would if they did not derive a convenience yield from holding Treasuries (propostion 2). Third, an increase in the supply of euro area government bonds induces a decline in equilibrium excess returns for Treasuries (3). I test by estimating 2SLS models on data from euro area banks and insurances. The first and third predictions can be tested directly against the sign of the first- and second-stage coefficients. On the contrary, the second prediction concerns an unobserved counterfactual, so I need to estimate the structural parameters γ^I and ψ to test it within the model.

3.1 Mapping the model to the data

In order to map regression coefficients to the structural parameters directly, in the following section I lay out a linearised version of the model. In addition to testing the predictions on the signs of derivatives, I use the restrictions implied by the linearised model to recover estimates for the structural parameters τ^B, τ^I and ψ. These estimates allow me to test the prediction of proposition 2 through model counterfactuals. Furthermore, they are informative on the extent to which cross-sector differences in sensitivity to excess returns are attributable to heterogeneity in risk aversion versus preference for non-monetary payoffs offered by US Treasuries.

Linearised model

I linearise the model by taking a first-order Taylor expansion of the expressions for b^B_{US} b^I_{US}, and ϕ around the following points for endogenous variables b^B_{US}, b^I_{US}, and $\mathbb{E}[R_{US} - R]$: [6]

$$\bar{b}^B_{US} = \frac{1}{2}$$
$$\bar{b}^I_{US} = \frac{1}{2}$$
$$\bar{\mathbb{E}}[R_{US} - R] = 0$$

[6] I use ϕ rather than $\mathbb{E}[R_{US} - R]$ because the empirical model isolates changes in the convenience yield component of excess returns, in keeping with the focus of the paper.

Combined with the budget constraints, the market clearing conditions, and the assumptions $W_0^B = W_0^I = 1$, these choice of symmetrical portfolio shares as approximation points implies $\bar{B} = 1$ and $\bar{B}_{US} = 1$.

The variance and covariance terms are taken as fixed parameters, calibrated to their sample counterpart in the estimation of structural parameters. Proposition 4summarises the three resulting equations:

Proposition 4 (Linearised model)**.** (i) *US Treasury portfolio share for banks*
$s_{US}^B = \frac{1}{\gamma^B \mathbb{V}[R_{US}-R]} \left(\mathbb{E}[R_{US} - R] - \gamma^B Cov[R_{US} - R, Y^B] \right)$.

(ii) *US Treasury portfolio share for insurances*
$s_{US}^I \approx \frac{\mathbb{E}[R_{US}-R]}{2\gamma^I \mathbb{V}[R_{US}-R]} \left(1 + \frac{\gamma^I Cov[R_{US}-R, Y^I]}{\sqrt{(\gamma^I Cov[R_{US}-R, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US}-R]}} \right)$.

(iii) *Equilibrium convenience yield*
$\phi \approx \frac{\psi \tau^I \mathbb{V}[R_{US}-R]}{\psi \tau^B \tau^I + 0.25 \mathbb{V}[R_{US}-R](\tau^B + \tau^I)} \left(1.5 + \frac{Cov[R_{US}-R, Y^B]}{\mathbb{V}[R_{US}-R]} + \left(\frac{\tau^B}{\tau^B + \tau^I} - 1 \right) B \right)$

Proof. In Appendix C.4. □

Note the use of the equality rather than approximation symbol for bank portfolio shares, as they are already exactly linear in excess returns.

The linearised model can be re-written as a set of three estimation equations with s_{US}^B, s_{US}^I, ϕ and $\mathbb{E}[R_{US} - R]$ as observables:

$$s_{US}^B = \beta^B \phi + \varepsilon^B, \tag{8}$$
$$s_{US}^I = \beta^I \phi + \varepsilon^I, \tag{9}$$
$$\phi = \pi B + \nu. \tag{10}$$

The parameters of the empirical model as a function of structural parameters are

$$\beta^B = \frac{\partial s_{US}^B}{\partial \phi} := \frac{\tau^B}{\mathbb{V}[R_{US} - R]} > 0, \tag{11}$$

$$\beta^I = \frac{\partial s_{US}^I}{\partial \phi} := \frac{\tau^I}{2\mathbb{V}[R_{US} - R]} \left(1 + \frac{Cov[R_{US} - R, Y^I]}{\tau^I \sqrt{\left(\frac{1}{\tau^I} Cov[R_{US} - R, Y^I]\right)^2 + \frac{4\psi \mathbb{V}[R_{US}-R]}{\tau^I}}} \right) > 0, \tag{12}$$

$$\pi = \frac{\partial \phi}{\partial B} := \frac{\psi \tau^I \mathbb{V}[R_{US} - R]}{\psi \tau^B \tau^I + 0.25 \mathbb{V}[R_{US} - R](\tau^B + \tau^I)} \left(\frac{\tau^B}{\tau^B + \tau^I} - 1 \right) < 0 \tag{13}$$

The error terms are

$$\varepsilon^B := \frac{\tau^B}{\mathbb{V}[R_{US}-R]}\left(RP - \gamma^B Cov[R_{US}-R, Y^I]\right),$$

$$\varepsilon^I := \frac{\tau^I RP}{2\mathbb{V}[R_{US}-R]}\left(1 + \frac{Cov[R_{US}-R, Y^I]}{\tau^I\sqrt{\left(\frac{1}{\tau^I}Cov[R_{US}-R, Y^I]\right)^2 + \frac{4\psi\mathbb{V}[R_{US}-R]}{\tau^I}}}\right),$$

$$\nu := \frac{\psi\tau^I\mathbb{V}[R_{US}-R]}{\psi\tau^B\tau^I + 0.25\mathbb{V}[R_{US}-R](\tau^B+\tau^I)}\left(1.5 + \frac{Cov[R_{US}-R, Y^B]}{\mathbb{V}[R_{US}-R]}\right)$$

Note that the error terms depend on the risk premium component of excess returns RP. If unaccounted for, this term would introduce bias in the estimation. The model assumes that risk premia are due solely to underlying fluctuations in exchange rates. However, I estimate the model on a sample including European sovereign bonds with non-negligible credit risk. In the empirical model, I control for time-varying risk premia by using a proxy for ϕ that adjusts for both forward rates, partialling out exchange rate risk, and CDS rates, partialling out sovereign risk. Furthermore, the model accounts for constant differences in risk premia across euro area sovereign bonds via country fixed effects.

I estimate β^B, β^I and ϕ via a 2SLS procedure with equation 10 as first stage, and equations 8 and 9 as second stages. I identify β^B and β^I using exogenous changes in B as an instrument for ϕ in both 8 and 9. Equation 10 is obtained by linearising ϕ as a function of B and b^B, expressing b^B as a function of b^B_{US}, and then writing out b^B_{US} as a function of ϕ and RP. As a result, the linear model implies that the estimation equation for ϕ does not depend on b^B_{US}. In other words, b^B_{US} depends on ϕ only through B, so the exclusion restriction for using the latter as an instrument for ϕ in a model of b^B_{US} is satisfied. Note that the model predicts the same reaction of portfolio shares to excess returns regardless of whehter the latter results from a change in the risk premium or the convenience yield component. Therefore, a valid instrument for ϕ identifies $\frac{\partial b^X_{US}}{\partial\phi} = \frac{\partial b^X_{US}}{\partial\mathbb{E}[R_{US}-R]}$, for $X = \{B, I\}$. I can then focus solely on ϕ in the empirical model, while resting assured that the restrictions on structural parameters implied by $\frac{\partial b^X_{US}}{\partial\mathbb{E}[R_{US}-R]}$ remain unchanged.

I can then exploit the restrictions imposed by estimated coefficients $\hat{\beta}^B$, $\hat{\beta}^I$, and $\hat{\pi}$ to back out estimates for structural parameters, as detailed in section 4.1.

3.2 Data

I estimate the equations outlined in the previous section on data from banks and insurances resident in the euro area, sourced from the publicly available Securities Holdings Statistics (SHS) dataset by the European Central Bank. The European setting provides an ideal setting to study the role of convenience yields for the demand of US Treasuries by different types of investors due to two features.

First, the available data allows us to observe the sovereign portfolios of different sectors, banks and insurances, operating under essentially the same regulatory regime. Existing literature on the demand for government bonds mainly relies on global data that does not provide a breakdown of government bond holdings by both country and sector (Tabova and Warnock, 2022, Eren et al., 2023, Fang et al., 2023). While cross-section differences in preferences are plausibly constant across jurisdictions, the regulatory regimes for banks relative to insurances might not be. Therefore, the cross-sector differences in the sensitivity of the demand for safe assets estimated in previous studies might confound differences in preferences and regulation. In the European Union, banks and insurances are subject to very similar rules concerning investment in sovereign debt. Both Article 351 of the Capital Requirement Regulation (*EU Regulation No 575/2013*), applying to banks; and Article 180 of the Solvency II regulation(*EU Regulation No 35/2013*), applying to insurances, assign, with almost identical language, a zero weight for capital requirements to bonds that either have a high sovereign rating (like US Treasuries) or are denominated in euros. Therefore, from the regulatory point of view, both banks and insurances are free to adjust the relative portfolio share of US and euro area sovereign bonds in response to returns without affecting their stock of risk-weighted assets relevant for capital buffers.[7] As a result, any observed discrepancy in the demand sensitivity to convenience yields across these sectors is plausibly attributable to preferences. Thanks to this regulatory design, I can zoom in on differences in preferences only, using the estimates of structural parameters to disentangle the role of risk aversion and convenience yields.

Furthermore, the use of global data for different investor classes would introduce complications in mapping the model to the data. The theoretical model in this paper analyses the simple choice between US and domesic. While euro-

[7]The only material difference in the regulatory treat . Exposure to US Treasuries, if unhedged, counts against regulatory limits for foreign exchange risk exposure. Faia et al. (2022) uses this divergence, together with a different regime of capital requirements between insurance companies and mutual funds, to motivate differences in demand elasticity for these two sectors and deviations from covered interest parity. In this paper, I abstract from limits to foreign exchange exposure, which would affect both banks and insurances equally.

denominated sovereign bonds are a natural choice of domestic asset when focusing on the eurozone, this would not be the case when using global data. [8]. Extending the exercise of this paper to global data would require either a more complex model of the whole sovereign portfolio with multiple assets, or the construction of a synthetic "domestic" asset for global foreign investors in US Treasuries from the data.

Second, the peculiar structure of purchases under the PSPP quantiative easing policy by the ECB generates exogenous variation in the relative supply of government bonds across euro area countries. Combined with the differences in convenience yields between sovereign bonds in the eurozone, documeted by Jiang et al. (2020), this policy intervention provides an ideal instrument to identify exogenous changes in the convenience yield component of US Treasury excess returns relative to euro area sovereign bonds. Therefore, I can estimate equations 8, 9 and 10 with a panel 2SLS strategy.

3.3 Identification strategy

According to the model, changes in the supply of euro area government bonds B are a valid and relevant instrument for the convenience yield term of excess returns in estimating the portfolio equations 8 and 9, because the latter depend on B only on excess returns. However, the very stylised partial equilibrium model does not take into account that debt supply is likely endogenous to the portfolio choice of financial intermediaries through general equilibrium effects. Even considering only changes in debt supply due to unconventional monetary policy is not enough to allay concerns of endogeneity, as these policies are adopted in response to highly endogenous macroeconomic development. This argument has particular bite for European banks, as they have been observed to load up on domestic government bonds in precisely the same turbulent macroeconomic conditions that motivate quantiative easing policies, either in a "gambling for resurrection" strategy (Acharya and Steffen, 2015) or due to "moral suasion" by their governments (De Marco and Macchiavelli, 2016, Ongena et al., 2019).

In order to obtain changes that are truly exogenous to investors' portfolio choice, I exploit the characteristics of the PSPP, implemented by the ECB from January 2015. The ECB bought government bonds issued by all countries with

[8] I use data on the portfolios of the aggregate banking and insurance sector in the eurozone, so the domestic asset is defined at the currency rather than country level. This approach is consistent with the assumption of risk premia stemming only from exchange rate fluctuations in the model.

a credit rating of at least BBB-, and with maturities from 2 to 30 years. [9] The purchases are apportioned according to a scheme that favoured a market-neutral approach. They are proportional to each country's capital key, and they mirror as closely as possible the maturity structure of outstanding bonds.

The capital key for each country is the equal-weighted average of its share of the euro zone's population and GDP. It is updated every five years, and whenever the membership of the European Union (EU) changes. In my sample, running from 2015 to 2022, the capital key changed twice: in 2019 due to a five-yearly update, and in 2020 due to the withdrawal of the United Kingdom from the EU. Since country size is plausibly independent of portfolio choice, and updates related to GDP are slow-moving, changes in capital key are likely exogenous. The other source of variation is the cross-country difference between the extant maturity structure of PSPP holdings, and that of the country's outstanding government bonds. Since this difference depends only on the governments' choice on the maturity of issuance and on the pre-existing PSPP term structure, it probably satisfies the exclusion restriction as well. Koijen et al. (2021) also uses PSPP purchases as an instrument for debt supply, but it relies on purchases predicted by the capital key rather than the actual amounts.

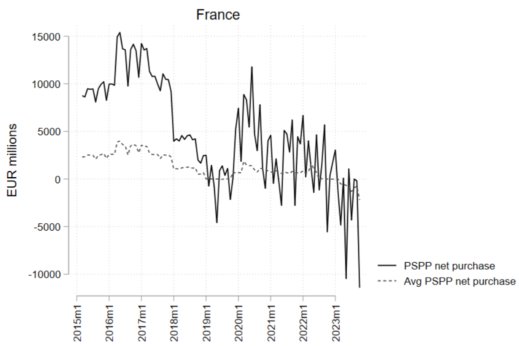
One step is missing to be able to claim PSPP holdings are a valid instrument for ϕ: while it can be argued that the *cross-sectional* variation in PSPP holdings is exogenous to investors' portfolio choice, this is not the case for the *time-series* dimension. Changes in the total purchases over time track the overall size of the quantiative easing programme, which is obviously correlated to investors' portfolios through macroeconomic fluctuations. To account for this issue, I use time fixed effects that soak up trends in the average amount of PSPP purchases across countries illustrates the intuition behind this strategy. The solid line represents purchases for each of the four largest euro area country, while the dashed line depicts average purchases across countries. By using time fixed effects, I rely only on the informtion contained in the differences between the grey and blue line for each country.

3.4 Estimation via two-stage least squares

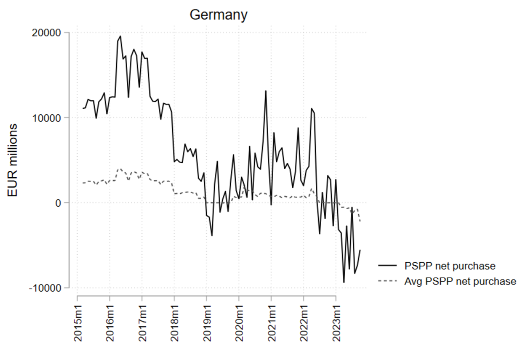
In this section, I lay out the estimating equations for the empirical model and report the results. The starting point is the set of three equations derived from

[9] The restriction on credit rating resulted in the exclusion of Greek bonds. In the context of this paper, the exclusion of Greece helps in isolating the convenience yield component, as it would likely be obscured by the extremely large risk premium on Greek government bonds.

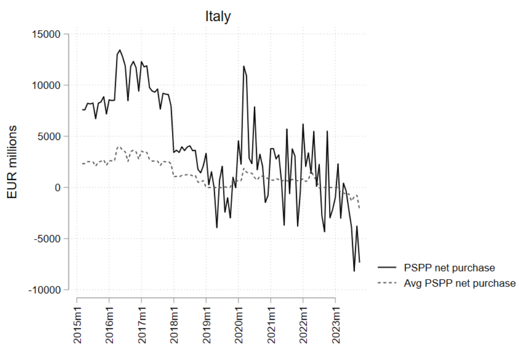
Figure 3. PSPP purchases



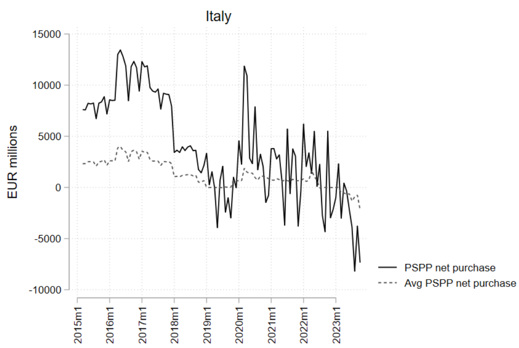
(a) France



(b) Germany



(c) Italy



(d) Spain

Monthly net purchases of sovereign debt under the Public Sector Purchase Programme, all maturities. The solid line depicts monthly net purchases for a country, the dashed line represents the cross-sectional average of monthly net purchases across all eligible countries over quarters. Source: European Central Bank

the linearised theoretical model. While they can be estimated directly as written above, I modify them to account for complications and nuances in the data that the model fails to capture. The baseline first-stage regression is

$$\phi_{i,t} = \alpha_i + \alpha_t - \pi PSPP_{i,t} + \lambda' \mathbf{V_{i,t}} + \kappa' \mathbf{W_{i,t}} + \nu_{i,t},$$

where $PSPP_{i,t}$ are PSPP holdings of country i government debt in quarter t, which proxy the exogenous component of B.[10] Note tht the coefficient π enters the

[10]Note that section 3.3 discusses the identification strategy in terms of PSPP purchases to build intuition, but here I use PSPP *holdings* because in the theoretical model excess returns

equation with a minus sign because an increase in PSPP holdings corresponds to a *decrease* in the amoun of country i's sovereign debt available to investors. The second-stage regression for $X = \{B, I\}$ is

$$s^X_{US,i,t} = \alpha^X_i + \alpha^X_t + \beta^X \phi_{i,t} + \delta^{X\prime}\mathbf{V_{i,t}} + \eta^{X\prime}\mathbf{W_{i,t}} + \varepsilon^X_{i,t}.$$

One difference from the theoretical model is due to the panel structure of the data. I rely on quarterly observations of sovereign holdings, so $s^X_{US,i,t}$ is the quarter t share of US Treasuries in a portfolio comprised of country i's government bond and US Treasuries for all euro area investors in sector X. Likewise, $\phi_{i,t}$ is the convenience yield portion of excess returns of US Treasuries with respect to country i's government bonds, on average for quarter t. The convenience yield term is a theoretical concept, so I need an observable proxy. I follow the approach in Du et al. (2018) and approximate the convenience yield as follows:

$$\phi_{i,t} = y_{US,t} - y_{i,t} + \rho_t - bs_{i,t} - l_{i,t}$$

where $y_{US,t}$ and $y_{i,t}$ are the yields of US and country i government bonds; ρ_t is the market-implied forward premium for the euro against the dollar; $bs_{i,t}$ with is the EUR/USD cross-currency basis swap, a measure of CIP deviations in inter-bank rates, and $l_{i,t} \coloneqq CDS_{US,t} - CDS_{i,t}$ is the difference in sovereign CDS rates between the US and country i. All components are averaged over the 1,2,3,5, and 10 year maturities and over quarter t. After correcting for fluctuations in exchange rates ($\rho_{i,t}$), differences in credit risk ($l_{i,t}$), and frictions in forex markets ($bs_{i,t}$), the residual difference in yields between US and country i's government bonds reflect only relative convenience yields.[11]

The regressions also include fixed effects at both the quarter (α_t) and country (α_i) level. As explained in section 3.3, time fixed effects aid the identification strategy by isolating cross-country differences in PSPP holdings. Furthermore, they control for any global determinants of the demand and supply of safe assets. Country fixed effects account for time-invariant featrues such as idiosyncratic country risk, which is part of the risk premium in the error terms derived in section 3.1. Controlling for country risk is particularly important for bonds issued by distressed sovereign such as Italy, Ireland, Portugal and Spain. Given the "bank-sovereign nexus", whereby investment in distressed sovereign bonds by domestic banks are often driven by political considerations (Andreeva and Vlassopoulos, 2016, Ongena et al., 2019, Saka, 2020), country-specific risk is especialy relevant in the model for

depend on the *level* of government debt supply.

[11] Du et al. (2018) defines the convenience yield as a premium for US Treasury, and defines it as $y_{US,t} - y_{i,t} - \rho_t + bs_{i,t} + l_{i,t}$. In this paper I define the convenience yield as a component of Treasury *excess returns*, so it has the opposite sign.

bank portfolios. Country fixed effects also take care of any potential cross-country pattern in the correlation and variance of excess returns, which enter the error term as shown in section 3.1.

I also augment the model with two sets of country-quarter level controls. The first set, $\mathbf{V_{i,t}}$ accounts for changes in the portfolio shares due to valuation effects. The theoretical model is cast in real terms, so in the data I need to control for valuation effects due to both bond prices and exchange rates, in order to isolate actual portfolio rebalancing. In the baseline specification with time and country fixed effects, $\mathbf{V_{i,t}}$ includes quartely changes in the all-maturity price index for country i's government bonds $\Delta BI_{US,t}$. Note that changes in the EUR/USD exchange rate and in the dollar price of US bonds, which also affect portfolio shares, are subsumed in the quarterly fixed effects. They are included in $\mathbf{V_{i,t}}$ for specifications whose fixed effect structure allow it.

The second set of controls $\mathbf{W_{i,t}}$ include CPI inflation, real GDP growth and the ratio of government debt to GDP for country i in quarter t. The first two variables are included to succinctly capture macroeconomic factors that affect investment in country i, which might be correlated with the maturity choice of government debt issuance, in turn driving variation in PSPP holdings. The latter captures both a time-varying factor of country risk, and possible concerns of residual discretionality in PSPP purchases tilted towards highly indebted countries.

I estimate the model on country-quarter observations from 2015 Q1 to 2022 Q1 for all eurozone countries except Greece, as it is excluded from the PSPP, and Estonia and Luxembourg, due to data availability.

First-stage regression

Table 1 displays results from the first-stage regression. The coefficient on $PSPP_{i,t}$ in the column 3, the preferred specification including both time and country fixed effects, reports a highly statistically significant increase of 0.38 percentage points in the convenience yield component of US Treasury excess returns in response to a one standard deviation increase in PSPP holdings, equivalent to about \$ 153 billion. The size of the coefficient is also significant, on the same order of magnitude of the 0.24 percentage point standard deviation of convenience yields.

The sign of the coefficient is consistent with the prediction of the model, as an increase in PSPP holdings corresponds to a decrease in the supply of country i government debt on the market. The increase in US Treasury excess returns through the convenience yield component echoes the findings of Jiang et al. (2021b) , which highlights the change in the relative supply of safe assets and convenience yields as an additional channel through which quantitative easing affects equilibrium interest rates.

Columns 1 and 2 show the results of models with no fixed effects, and with time fixed effects only, respectively. The inclusion of time fixed effects is crucial in my identification strategy as it allows to single out exogenous variation in relative debt supply that is independent of the ECB's unconventional monetary polcy stance. Column 2 shows how time fixed effects raise the F stat to 12.42, implying a stronger instrument as well as an arguably more valid one. Column 3 shows that the inclusion of country fixed effects does reduce the F statistic. Therefore, there is a tradeoff between having a strong instrument and controlling for important country-specific factors such as credit risk and the sovereign-bank nexus.

Table 1. First-stage regression

	(1)	(2)	(3)
$PSPP_{i,t}$	0.21***	0.22***	0.38***
	(0.07)	(0.06)	(0.14)
$\Delta BI_{i,t}$	-0.03	-0.12	-0.08
	(0.06)	(0.09)	(0.07)
$\Delta BI_{US,t}$	0.03		
	(0.08)		
$\Delta e_t^{EUR/USD}$	-0.04		
	(0.05)		
N	309	309	309
F stat	8.24	12.42	7.13
Macro controls	Yes	Yes	Yes
Time fixed effects	No	Yes	Yes
Country fixed effects	No	No	Yes

Coefficients from regression model $\phi_{i,t} = \alpha_i + \alpha_t + \pi PSPP_{i,t} + \lambda' \mathbf{V_{i,t}} + \kappa' \mathbf{W_{i,t}} + \nu_{i,t}$. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Second-stage regressions

Tables 2 and 3 report the estimation results for banks and insurances, respectively.

The preferred 2SLS specification with time and country fixed effects shows an increase of the US Treasury portfolio shares in response to a higher convenience yield component of excess returns. The coefficient is statistically significant at the 1% level for both banks and insurances. Therefore, the prediction of a positive reaction to excess returns is verified for both types of investors.

The comparison of coefficients on $\phi_{i,t}$ estimated via OLS (columns 1 to 3) and 2SLS (columns 4 to 6) suggests that the instrumental variable strategy purges the coefficients from the bias due to the endogeneity of supply and demand. The coefficient β represents the sensitivity of portfolio shares, an equilibrium quantity, to a component of excess returns, tightly connected to equilibrium asset prices. Therefore, the observed price-quantity data points are likely driven by both demand and supply shocks. The former introduce a negative correlation between $s^X_{US,i,t}$ and $\phi_{i,t}$, while the latter a positive correlation. Therefore, failing to isolate supply shock would result in a bias toward zero for β. PSPP-induced exogenous changes in B act as a supply shock, so by using them as instrument for $\phi_{i,t}$ the slope of the demand curve can be recovered. The larger and more significant coefficients across the board in the 2SLS models indicate that this strategy is indeed successful.

In order to understand what global variables are accounted for by time fixed effects, models in columns 1 and 4 replace them with the VIX, a key determinant of the demand for safe assets (Miranda-Agrippino and Rey, 2022), and with the debt/GDP ratio in the US to control for the supply of US Treasuries. The estimated β in columns 1 and 4 are very similar to those in columns 2 and 5, which use the same specification but replace global controls with quarterly fixed effects. Therefore, time fixed effects appear to capture well the role of global drivers of the demand and supply of safe assets. The lack of time fixed effects also allows me to augment the vector of valuation effect controls $\mathbf{V_{i,t}}$ with changes in the exchange rate and US bond prices, which vary only in the time dimension. However, none of the valuation effects are statistically significant even at the 10% level, possibly because of relatively small quarter-on-quarter variation.

The models estimated in columns 2 and 5 include time fixed effects, but not country fixed effects. Comparing them to the coefficients in columns 3 and 6 reveals the imporance of controlling for country-specific characteristics, especially for banks, as argued in the previous section. In fact, the eclusion of country fixed effects in the bank regressions results in very large negative coefficients, that do not seem particularly credible due to both their sign and magnitude. The lower sensitivity of the insurance models to country fixed effects corroborates the hypothesis that the political economy factors affecting investment of European banks in distressed sovereign bonds contribute to biasing the estimates. However, it is still important to include country fixed effects in the model for insurances. The reasons lie both in conssitency with the estimates for banks, and to account for the time-invariant portion of country-specific credit risk, which the theoretical model abstracts away.

Table 2. Second-stage regression for banks

	OLS			2SLS		
	(1)	(2)	(3)	(4)	(5)	(6)
$\phi_{i,t}$	-0.81	-0.89	0.65***	-68.19***	-75.26***	6.93***
	(0.85)	(0.98)	(0.21)	(24.01)	(21.09)	(2.50)
$\Delta BI_{i,t}$	0.05	0.98	-0.20	-2.65	-8.87	0.32
	(0.83)	(1.38)	(0.16)	(4.47)	(7.31)	(0.46)
$\Delta BI_{US,t}$	0.18			3.43		
	(1.09)			(6.03)		
$\Delta e_t^{EUR/USD}$	0.19			-3.32		
	(0.75)			(4.10)		
N	309	309	309	309	309	309
Macro controls	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	Yes	Yes	No	Yes	Yes
Country fixed effects	No	No	Yes	No	No	Yes
Underid test p-value				0.01	0.00	0.01
Weak id test stat				8.24	12.42	7.13

Coefficients from regression model $s^B_{US,i,t} = \alpha^B_i + \alpha^B_t + \beta^B \phi_{i,t} + \delta^{\mathbf{B}'}\mathbf{V_{i,t}} + \eta^{\mathbf{B}'}\mathbf{W_{i,t}} + \varepsilon^B_{i,t}$ estimated via OLS, or via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The underidentification test uses the Kleibergen and Paap (2006) LM statistic. The weak identification test uses the Kleibergen and Paap (2006) Wald F statistic. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Figure 4 compares the coefficients on $\phi_{i,t}$ from column 6 of the models for banks and insurances. The US Treasury portfolio share of banks increases by 6.93 percentage points in response to a one percentage point increase in the convenience yield component of excess returns, while the portfolio share of insurances increases by 2.31 percentage points. The coefficients are statistically different from each other at the 10% significance level. The muted reaction of insurances' portfolio shares to excess returns is consistent with results in the literature (Timmer, 2018, Eren et al., 2023, Koijen et al., 2021), and it suggests there is enough of a discrepancy in behaviour between the two sectors that more than risk aversion might be at play. However, it is not enough to test proposition 2 (ii), as it concerns a counterfactual on the behaviour of convenience yield investors. In order to

Table 3. Second-stage regression for insurances

	OLS			2SLS		
	(1)	(2)	(3)	(4)	(5)	(6)
$\phi_{i,t}$	0.44**	0.51*	0.03	12.74***	14.69***	2.31***
	(0.22)	(0.27)	(0.02)	(4.61)	(4.12)	(0.87)
$\Delta BI_{i,t}$	0.76	2.12	-0.16	-2.01	-7.94	-1.21
	(0.94)	(1.63)	(0.14)	(4.17)	(7.13)	(0.80)
$\Delta BI_{US,t}$	-0.03			3.87		
	(1.26)			(5.68)		
$\Delta e_t^{EUR/USD}$	0.43			-3.14		
	(0.81)			(3.88)		
N	307	307	307	307	307	307
Macro controls	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	Yes	Yes	No	Yes	Yes
Country fixed effects	No	No	Yes	No	No	Yes
Underid test p-value				0.01	0.00	0.01
Weak id test stat				8.01	13.23	7.30

Coefficients from regression model $s^I_{US,i,t} = \alpha^I_i + \alpha^I_t + \beta^I \phi_{i,t} + \delta^{\mathbf{I}'}\mathbf{V_{i,t}} + \eta^{\mathbf{I}'}\mathbf{W_{i,t}} + \varepsilon^I_{i,t}$ estimated via OLS, or via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The underidentification test uses the Kleibergen and Paap (2006) LM statistic. The weak identification test uses the Kleibergen and Paap (2006) Wald F statistic. Robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

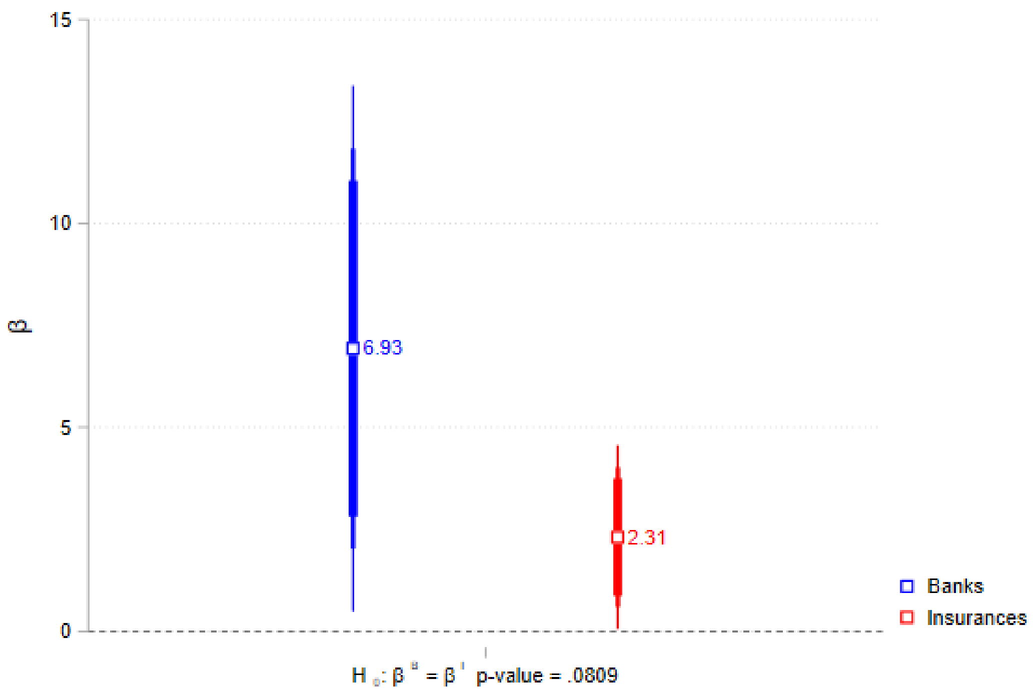
quantify the relative importance of risk aversion and preference for non monetary payoffs it is necessary to recover the structural parameters γ^B, γ^B and ψ from the estimates of the empirical model. I turn to this task in the next section.

4 Structural parameters and model counterfactuals

4.1 Recovery of structural parameters

After obtaining the coefficient estimates $\hat{\beta}^B$, $\hat{\beta}^I$ and $-\hat{\pi}$ from the 2SLS model, I can back out values for the structural parameters by solving the system of equations 11, 12 and 13 for $\tau^B := \frac{1}{\gamma^B}$, $\tau^I := \frac{1}{\gamma^I}$ and ψ. To do so, I replace β^B, β^I and π with their estimates, and proxy $\mathbb{V}[R_{US} - R]$ and $Cov[R_{US} - R, Y^I]$ with their

Figure 4. Comparison of β for banks and insurances



The blue-bordered dot represnets the coefficient β^B, and the red-bordered dot represents the coefficients β^I. Both coefficients are estimated via 2SLS in the model $s^I_{US,i,t} = \alpha^I_i + \alpha^I_t + \beta^I \phi_{i,t} + \delta^{\mathbf{I}'}\mathbf{V_{i,t}} + \eta^{\mathbf{I}'}\mathbf{W_{i,t}} + \varepsilon^I_{i,t}$ via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The bars around the dots represent confidence intervals at the 90%, 95% and 99% levels, in decreasing order of thickness. The p-value on the hypothesis $H_0 : \beta^B = \beta^I$ is calculated using the Clogg et al. (1995) method, which assumes that the coefficients are independent.

sample counterparts. [12]

Table 4 reports the distribution of recovered structural parameters. Insurances display a lower risk tolerance, with a mean of 1.56 compared to 15.4 for banks. The size of the parameter ψ reveals that convenience yields play a non-negligible role in insurers' preferences. The mean value of 0.11 corresponds to a weight of 10% in their objective function. The relative size of the parameters suggests a stronger role for risk aversion, but alone it is not enough to pin down the relative importance of the two components. In the next section, I perform counterfactual experiments on

[12] Appendix F details the procedure to solve the system and simulate the distribution of structural parameters.

Table 4. Structural parameters

Structural parameter	Mean	95% CI lower bound	95% CI upper bound
τ^B	15.4	6.2	24.63
τ^I	1.56	-12.96	14.67
ψ	0.11	0.0	0.53

Confidence intervals are obtained by drawing 100000 times from the joint asymptotic distribution of parameters in the empirical model, solving for structural parameters for each joint draw, and computing the 5th and 95th percentiles of the simulated distribution.

ψ and τ^I to investigate how parameter values translate into the relative strength of risk aversion and convenience yields in determining both portfolio choice and equilibrium interest rates.

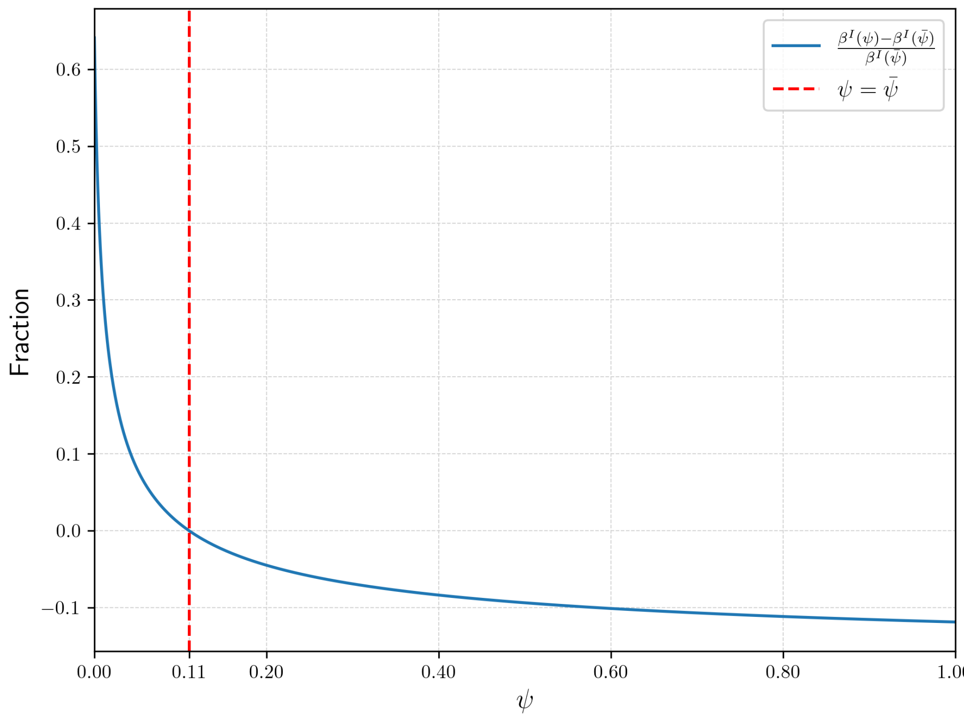
4.2 Model counterfactuals

First, I perform two distinct but related exercises to quantify the importance of risk aversion versus convenience yields in explaining the magnitude of insurers' portfolio share sensitivity, and its difference from the sensitivity of banks. Then, I investigate the effect of convenience yield investors on the reaction of equilibrium excess returns to changes in debt supply.

Portfolio share sensitivity as a function of ψ

Proposition 2 (ii) claims that the reaction of to excess returns is smaller than it would be in the absence of non-monetary payoffs in the objective function, that is $\psi = 0$. To test it, I calculate the counterfactual values of β^I as a function of ψ fixing τ^B and τ^I at their means. To get a sense of the magnitude of the effect of ψ on the sensitivity to excess returns, I subtract $\beta^I(\bar{\psi})$, the value of β^I at the mean for ψ, and divide the difference by it. Figure 5 plots this function against ψ (blue line). For values lower than the mean $\bar{\psi}$ (red dashed line), the function is positive, meaning that the corresponding β^I is larger and hence that proposition 2 (ii) is verified. The sensitivity coefficient β^I at $\psi = 0$, in the absence of prefernces for non-monetary payoffs, is 60% larger than the coefficient corresponding to the average weight of the convenience yield term, $\bar{\psi}$. Therefore, the estimated parameters imply that the non-monetary payoff component of preferences has a sizeable impact on the yield sensitivity of insurances' US Treasury share.

Figure 5. Percentage change in β^I as a function of ψ

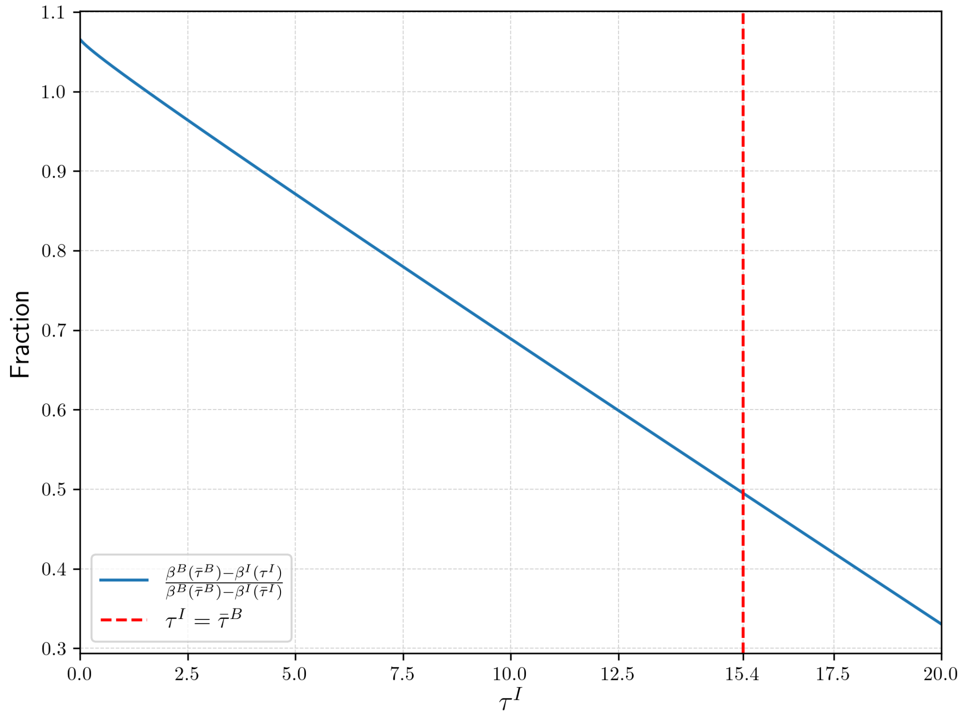


$\beta^I(\psi)$ is calculated using the means of parameters τ^B and τ^I drawn from the simulated distribution, letting ψ vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\psi})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of ψ from the simulated distribution, defined as $\bar{\psi}$.

Portfolio share sensitivity as a function of τ^I

This exercise is aimed at quantifying the percentage of the difference in portfolio share sensitivity between banks and insurers that can be attributed to convenience yields versus risk aversion. I compute the difference in coefficients $\beta^B - \beta^I$ as a function of τ^I, which influences only β^I. I then divide it by the same difference evaluated at the mean for both parameters, , to obtain a fraction. Figure 6 plots this function against τ^I (blue line). The function is decreasing because β^I is increasing in τ^I: all else equal, a higher risk tolerance translates into a stronger reaction to excess returns. The value at $\tau^I = \bar{\tau^B}$ (red dashed line) is of particular interest. By equalising the risk tolerance of insurance and banks, any residual difference in β between the two sectors is attributable to the preference for non-monetary payoffs. According to this measure, convenience yields can explain 50% of the difference in β, so it also plays a sizeable role in the observed cross-sector heterogeneity of reactions to excess returns.

Figure 6. Percentage change in $\beta^B - \beta^I$ as a function of ψ



$\beta^I(\tau^I)$ is calculated using the means of parameters τ^B and ψ drawn from the simulated distribution, letting τ^I vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\tau^I})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of τ^I from the simulated distribution, defined as $\bar{\tau^I}$. $\beta^B(\tau^B)$ is calculated using the mean of parameter τ^B drawn from the simulated distribution and the calibrated value for and $\mathbb{V}[R_{US} - R]$ displayed in Table A5.

Decomposition of changes in excess return

The previous two experiments demostrated that the preference for non-monetary payoffs of Treasuries displayed by insurances has a quantitatively strong effect on portfolio choice. However, the presence of "convenience yield" investors affects equilibrium excess returns as well . As it can be seen by $\frac{\partial \phi}{\partial B} < 0$, if at least some investors value US Treasuries for their non-monetary payoffs, an increase in the equilibrium supply of euro area government bond will result in a stronger decline in Treasury excess returns, as "convenience yield investors" are willing to forgo a larger portion of returns for the now relatively scarcer US Treasuries. An important policy implication is a potentially stronger transmission channel for quantiative easing policies.

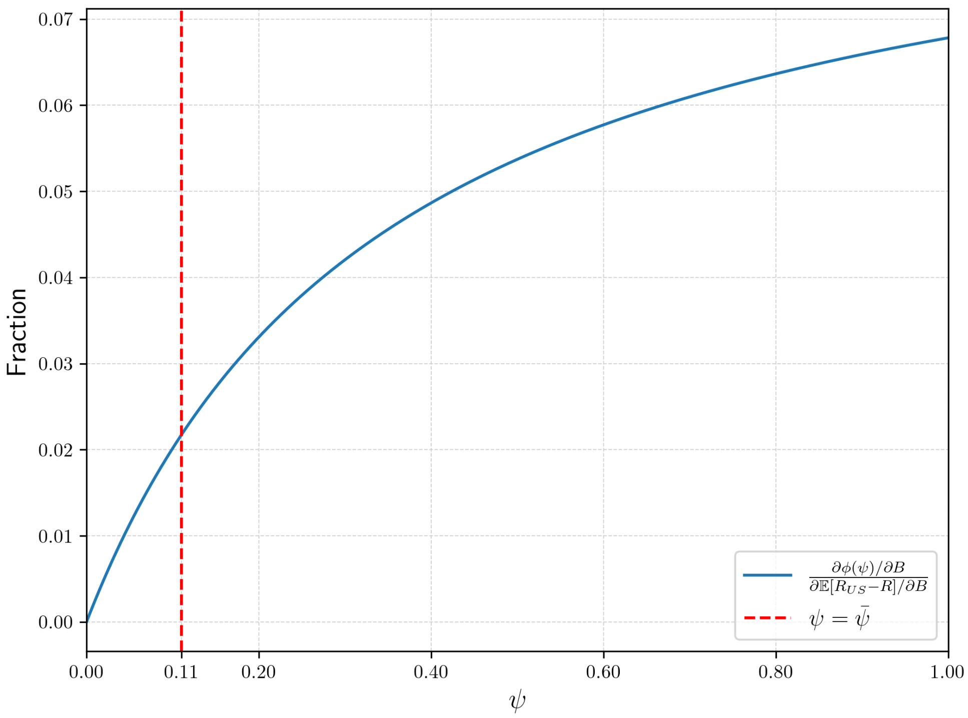
The recovered structural parameters pin down $\partial \mathbb{E}[R_{US} - R]/\partial B$, the change in equilibrium excess returns to an increase in B, through the linearised version of equations i and ii. Since excess returns are the simple sum of risk premium (RP) and convenience yield (ϕ) terms, I can use the decomposition $\partial \mathbb{E}[R_{US} - R]/\partial B = \partial RP/\partial B + \partial \phi/\partial B$ to compute the fraction attributed to the convenience yield as $\frac{\partial \phi/\partial B}{\partial \mathbb{E}[R_{US}-R]/\partial B}$. Therefore, I can quantify through the lense of the model the contribution of the convenience yield component on the total reaction of excess returns to changes in B, due for example to quantiative easing policies.

Figure 7 plots the percentage contribution of ϕ to changes in excess returns as a function of ψ. Intuitively, the function is increasing: the larger the weight of non-monetary payoffs, the larger their contribution to changes in excess returns, as convenience yield investor are willing to absorb a larger amount of Treasuries without requiring a higher yield. The figure shows that the convenience yield component is very small across the whole distribution of the estimated ψ, with a value of only 2% at the mean (red dashed line). However, this decomposition is specific to both the model and the European context of the data, so it does not necessary clash with the conclusion of other papers that find an important quantitative role for the convenience yield channel of unconventional monetary policy (Jiang et al., 2021b). In particular, the reliance on data for only two types of investors might understate the reaction of equilibrium excess returns.

5 Conclusion

This paper shows that the preference for the non-monetary payoffs of US Treasuries by certain classes of investors, such as insurances, plays a key role in explaining the observed differences in demand sensitivity across sectors. Risk aversion plays an equally important role in portfolio choice, but alone it cannot reconcile the joint observation of positive US Treasury portfolio shares for insurances, negative excess returns of Treasuries with respect to eurozone government bonds, and a positive correlation between Treasury excess returns and the income of insurances.

Figure 7. Proportion of $\partial\mathbb{E}[R_{US}-R]/\partial B$ explained by the convenience yield component



$\beta^I(\tau^I)$ is calculated using the means of parameters τ^B and ψ drawn from the simulated distribution, letting τ^I vary, and using the calibrated values for $Cov[R_{US}-R, Y^I]$ and $\mathbb{V}[R_{US}-R]$ displayed in Table A5. $\beta^I(\bar{\tau^I})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of τ^I from the simulated distribution, defined as $\bar{\tau^I}$. $\beta^B(\tau^B)$ is calculated using the mean of parameter τ^B drawn from the simulated distribution and the calibrated value for and $\mathbb{V}[R_{US}-R]$ displayed in Table A5.

Thanks to the estimation of structral parameters in investors preferences, the relative importance of convenience yields and risk aversion can be disentangled. Both a high risk aversion and a preference for non-monetary payoffs imply a lower reaction of portfolio shares to excess returns, so a structural approach is necessary to quantify their impact separately.

The model implies that, absent convenience yields, the demand of insurances would be 60% more sensitive to yields, resulting in a much less reliable absorption of additional US government debt by foreign investors.

Perhaps surprisingly, the sizeable impact of convenience yields on the yield sensitivity of portfolios is not matched by an equally large effect on equilibrium rates. The decomposition in the model shows that the convenience yield term accounts for only 2% of the reduction in Treasury excess returns after an increase in eurozone government debt. The policy implications are twofold. First, the safe asset channel of quantiative easing might not be very effective in reducing interest rates. Second, the dominance of the risk premium component suggests that reliance on yield-insensitive demand by foreigners to absorb additional sovereign debt at a low rate might not be a sustainable strategy for funding persistent US government deficits.

While the quantitative conclusion of this paper are by costruction limited to the context of European banks and insurance companies, they nevertheless offer insights for further research on the nuances of the foreign demand for Treasuries by different sectors in a global perspective.

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A Data sources

Table A1. Data sources

Data	**Source**
Government bond holdings of eurozone banks and insurances	ECB Securities Holdings Statistics
Income of eurozone banks	ECB Consolidated Banking database
Income of eurozone insurances	EIOPA Insurance Statistics
Government debt purchases and holdings under PSPP	ECB
Government bond indices and yields	Refinitiv Eikon
Spot and forward exchange rates	Refinitiv Eikon
EUR/USD cross-currency basis swap	Refinitiv Eikon
CDS rates	Refinitiv Eikon
Amount of government debt outstanding	Bank for International Settlements
Real GDP growth	OECD
CPI inflation	ECB and IMF
Debt/GDP ratio	Eurostat and Federal Reserve Economic Data (FRED)

B Summary statistics

Table A2. Summary statistics

	N	Mean	SD	Min	P25	P50	P75	Max
A. Portfolio shares								
$s_{i,t}^{US,MFI}$	442	71.2	26.8	14.8	50.9	80.9	95.3	99.1
$s_{i,t}^{US,ICPF}$	425	48.96	33.19	2.59	16.71	46.82	80.59	98.83
B. Financial variables								
$\mathbb{E}[R_{US} - R]$	463	-0.26	0.64	-5.21	-0.49	-0.16	0.08	0.77
$\phi_{i,t}$	407	0.24	0.23	-1.27	0.15	0.24	0.35	1.81
$\Delta BI_{i,t}$	371	0.20	4.39	-10.69	-1.89	0.42	3.10	14.09
$\Delta BI_{US,t}$	371	0.5	2.4	-4.7	-0.5	0.6	1.5	7.2
$\Delta e_t^{EUR/USD}$	463	0.4	3.9	-5.6	-1.8	0.1	2.5	13.7
C. Macroeconomic variables								
$Debt/GDP_{i,t}$	372	87.1	30.0	36.3	62.3	83.3	108.4	158.9
$\Delta CPI_{i,t}$	372	1.4	1.9	-2.2	0.2	1.1	2.0	11.7
$\Delta GDP_{i,t}$	343	0.6	3.5	-17.6	0.2	0.5	0.9	21.4

Summary statistics calculated on the data in which observations for PSPP holdings are non-empty. All variables in percentage points.

C Proofs

C.1 Proposition 1: optimal portfolio share for insurances

Proof. Proposition 1 (i): substitute the constraints in the objective function to recast the problem with b^I_{US} as a choice variable. Recall that $b^I_{US} = s^I_{US}$ for $W^I_0 = 1$, so the following statements hold equally for s^I_{US}. Take the first-order condition for b^I_{US} to obtain the following quadratic equation:

$$\gamma^I \mathbb{V}[R_{US} - R](b^I_{US})^2 - \left(\mathbb{E}[R_{US} - R] - \gamma^I Cov[R_{US} - R, Y^I]\right) b^I_{US} - \psi = 0.$$

The two solutions are

$$b^I_{US,1} := \frac{\mathbb{E}[R_{US} - R] - \gamma^I Cov[R_{US} - R, Y^I]}{2\gamma^I \mathbb{V}[R_{US} - R]} - \frac{\sqrt{(-\mathbb{E}[R_{US} - R] + \gamma^I Cov[R_{US} - R, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US} - R]}}{2\gamma^I \mathbb{V}[R_{US} - R]}$$

and

$$b^I_{US,2} := \frac{\mathbb{E}[R_{US}-R] - \gamma^I Cov[R_{US}-R, Y^I]}{2\gamma^I \mathbb{V}[R_{US}-R]}$$
$$+\frac{\sqrt{(-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I])^2 + 4\gamma^I\psi\mathbb{V}[R_{US}-R]}}{2\gamma^I \mathbb{V}[R_{US}-R]}.$$

To select a solution, consider the conditions for $b^I_{US,1}, b^I_{US,2} > 0$, for for $\gamma^I > 0$, $\psi > 0$, and $\mathbb{V}[R_{US}-R] > 0$.

$$\begin{aligned} b^I_{US,1} > 0 \iff & \quad \mathbb{E}[R_{US}-R] - \gamma^I Cov[R_{US}-R, Y^I] \\ & > \sqrt{(-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I])^2 + 4\gamma^I\psi\mathbb{V}[R_{US}-R]} \\ \iff & \quad 4\gamma^I\psi\mathbb{V}[R_{US}-R] < 0 \end{aligned}$$

There is no solution for $\gamma^I > 0$, $\psi > 0$, and $\mathbb{V}[R_{US}-R] > 0$. Therefore, I choose the solution $b^I_{US} = b^I_{US,2}$, and derive the conditions for $b^I_{US,2} \in \mathbb{R}^+$ in the next proof.

Proposition 1 (ii): start with the conditions for $b^I_{US} \in \mathbb{R}$:

$$b^I_{US} \in \mathbb{R} \iff \left(\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I]\right)^2 - 4\gamma^I\psi\mathbb{V}[R_{US}-R] > 0.$$

It is immediate to see that it always holds for $\gamma^I > 0$, $\psi > 0$, and $\mathbb{V}[R_{US}-R] > 0$. Now consider the condition for $b^I_{US} > 0$.

$$\begin{aligned} b^I_{US} > 0 \iff & \quad -\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I] \\ & < \sqrt{(-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I])^2 + 4\gamma^I\psi\mathbb{V}[R_{US}-R]} \\ \iff & \quad 4\gamma^I\psi\mathbb{V}[R_{US}-R] > 0 \end{aligned}$$

The condition also always holds for $\gamma^I > 0$, $\psi > 0$, and $\mathbb{V}[R_{US}-R] > 0$. □

C.2 Proposition 2: sensitivity of insurances to excess returns

Proof. Proposition 2 (i): it follows immediately from differentiating b^I_{US} with respect to $\mathbb{E}[R_{US}-R]$. □

Proposition 2 (ii):

Proof.

$$\frac{\partial b^I_{US}}{\partial \mathbb{E}[R_{US}-R]} > 0 \iff$$
$$-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}, Y^I] > -\sqrt{(-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US}-R]}$$

This condition holds for $-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}, Y^I] > 0$, with $\gamma^I > 0$ and $\mathbb{V}[R_{US}-R] > 0$.

$$\frac{\partial b^I_{US}}{\partial \mathbb{E}[R_{US}-R]} < \frac{1}{\gamma^I \mathbb{V}[R_{US}-R]} \iff$$
$$\frac{1}{2}\frac{1}{\gamma^I \mathbb{V}[R_{US}-R]}\left(1 - \frac{-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I]}{\sqrt{(-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}-R, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US}-R]}}\right)$$
$$< \frac{1}{\gamma^I \mathbb{V}[R_{US}-R]} \iff$$
$$-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}, Y^I] < \sqrt{(-\mathbb{E}[R_{US}-R] + \gamma^I Cov[R_{US}, Y^I])^2 + 4\gamma^I \psi \mathbb{V}[R_{US}-R]}$$
$$\iff 4\gamma^I \psi \mathbb{V}[R_{US}-R] > 0$$

This condition always holds for $\gamma^I > 0$ and $\mathbb{V}[R_{US}-R] > 0$. □

C.3 Proposition 3: reaction of excess returns to euro area debt supply

Proof. Proposition 3 (i): it follows immediately from differentiating $RP := \frac{\mathbb{V}[R_{US}-R]}{\tau^B+\tau^I}(W_0^I + W_0^B - B) + \frac{Cov[R_{US}-R, Y^B] + Cov[R_{US}-R, Y^I]}{\tau^B+\tau^I}$ with respect to B. □

Proof. Proposition 3 (ii): differentiate $\phi := -\frac{\tau^I}{\tau^B+\tau^I}\frac{\psi}{W_0^I - B + b^B}$ with respect to B.

$$\frac{\partial \phi}{\partial B} = \frac{\tau^I}{\tau^B+\tau^I}\frac{\psi}{(W_0^I - B + b^B)^2}\left(-1 + \frac{\partial b^B}{\partial B}\right).$$

Since b^B depends on B only through equilibrium excess returns, apply the chain rule to write $\frac{\partial b^B}{\partial B} = \frac{\partial b^B}{\partial \mathbb{E}[R_{US}-R]}\frac{\partial \mathbb{E}[R_{US}-R]}{\partial B}$ and obtain the following:

$$\frac{\partial \phi}{\partial B} = \frac{\tau^I}{\tau^B+\tau^I}\frac{\psi}{(W_0^I - B + b^B)^2}\left(-1 + \frac{\partial b^B}{\partial \mathbb{E}[R_{US}-R]}\frac{\partial \mathbb{E}[R_{US}-R]}{\partial B}\right).$$

Using the budget constraint of banks $b^B + b^B_{US}$, re-write the derivative as a function of b^B_{US}.

$$
\begin{aligned}
\frac{\partial \phi}{\partial B} &= \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(W_0^I - B + b^B)^2} \left(-1 - \frac{\partial b^B_{US}}{\partial \mathbb{E}[R_{US} - R]} \frac{\partial \mathbb{E}[R_{US} - R]}{\partial B} \right) \\
&= \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(W_0^I - B + b^B)^2} \left(-1 - \frac{\partial b^B_{US}}{\partial \mathbb{E}[R_{US} - R]} \left(\frac{\partial RP}{\partial B} + \frac{\partial \phi}{\partial B} \right) \right) \\
&= \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(W_0^I - B + b^B)^2} \left(-1 - \frac{\tau^B}{\mathbb{V}[R_{US} - R]} \left(-\frac{\mathbb{V}[R_{US} - R]}{\tau^B + \tau^I} + \frac{\partial \phi}{\partial B} \right) \right)
\end{aligned}
$$

Collecting the $\frac{\partial \phi}{\partial B}$ terms,

$$
\frac{\partial \phi}{\partial B} = \frac{\tau^I \mathbb{V}[R_{US} - R]}{(\tau^B + \tau^I)\mathbb{V}[R_{US} - R](W_0^I - B + b^B)^2 + \tau^B \tau^I \psi} \left(\frac{\tau^B}{\tau^B + \tau^I} - 1 \right)
$$

For $\tau^B > 0$ $\tau^I > 0$, $\psi > 0$, and $\mathbb{V}[R_{US} - R]$, $\frac{\tau^I \mathbb{V}[R_{US}-R]}{(\tau^B+\tau^I)\mathbb{V}[R_{US}-R](W_0^I-B+b^B)^2+\tau^B\tau^I\psi} > 0$ and $\left(\frac{\tau^B}{\tau^B+\tau^I} - 1 \right) < 0$. Therefore, $\frac{\partial \phi}{\partial B} < 0$.

□

C.4 Proposition 4: linearised model

Proof. Proposition 4 (ii): take a first-order Taylor expansion of equation 3 around approximation point $\bar{\mathbb{E}}[R_{US} - R] = 0$. □

Proof. Proposition 4 (iii): take a first-order Taylor expansion of the ϕ term in equation 5 around approximation point $\bar{B}$, $\bar{b}^B$.

$$
\phi \approx \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(\bar{B} - W_0^I - \bar{b}^B)^2} (b^B - \bar{b}^B) - \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(\bar{B} - W_0^I - \bar{b}^B)^2} (B - \bar{B}).
$$

Writing b^B as a function of b^B_{US} from the banks' budget constraint,

$$
\phi \approx \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(\bar{B} - W_0^I - \bar{b}^B)^2} (W_0^B - b^B_{US} - \bar{b}^B) - \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(\bar{B} - W_0^I - \bar{b}^B)^2} (B - \bar{B}).
$$

Substituting out b^B_{US} from equation 1,

$$\phi \approx \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(\bar{B} - W_0^I - \bar{b}^B)^2} \left(W_0^B - \left(\frac{\tau^B}{\mathbb{V}[R_{US} - R]} (RP + \phi) - \frac{Cov[R_{US} - R, Y^B]}{\mathbb{V}[R_{US} - R]} \right) - \bar{b}^B \right) - \frac{\tau^I}{\tau^B + \tau^I} \frac{\psi}{(\bar{B} - WI_0 - \bar{b}^B)^2} (B - \bar{B}).$$

Collecting ϕ terms and substituting in the expression for RP,

$$\phi \approx \frac{\psi \tau^I \mathbb{V}[R_{US} - R]}{\psi \tau^B \tau^I + (\tau^B + \tau^I) \mathbb{V}[R_{US} - R] (\bar{B} - W_0^I - \bar{b}^B)^2} \left(\bar{B} + W_0^B - \bar{b}^B + \frac{Cov[R_{US} - R, Y^B]}{\mathbb{V}[R_{US} - R]} + \left(\frac{\tau^B}{\tau^B + \tau^I} - 1 \right) B \right)$$

For $W_0^B = W_0^I = 1$ and at the approximation point $\bar{b}^B = 0.5$, $\bar{B} = 1$, the expression reduces to

$$\phi \approx \frac{\psi \tau^I \mathbb{V}[R_{US} - R]}{\psi \tau^B \tau^I + 0.25 \mathbb{V}[R_{US} - R](\tau^B + \tau^I)} \left(1.5 + \frac{Cov[R_{US} - R, Y^B]}{\mathbb{V}[R_{US} - R]} + \left(\frac{\tau^B}{\tau^B + \tau^I} - 1 \right) B \right).$$

□

D Sensitivity of US Treasury demand along the yield curve

In this section, I investigate how the difference in the Treasury portfolio share sensitivity between banks and insurances changes along the yield curve. The special safety and liquidity features of US Treasuries vary according to maturity. Du et al. (2018) finds supporting evidence in the term structure of the Treasury premium, which is slightly higher at the 3-month and 1-year maturity. Theoretically, convenience yields are associated with "moneyness", so we can expect short-term securities to command a higher premium (Bansal et al., 2010, Nagel, 2016). On the other hand, duration matching of long-term liabilities with safe and liquid assets is a driving force between the special demand of insurances for US Treasuries (Greenwood and Vayanos, 2010, Greenwood and Vissing-Jorgensen, 2018). Therefore, this sector could plausibly derive a higher non-monetary payoff from long-term Treasuries instead. The exercise in this section can shed light on which of these two forces dominates in the context of European investors.

I run the baseline regressions in the preferred 2SLS specification with time and country fixed effects separately for short- and long-maturity. Consistently with the breakdown available in the SHS data, I define short-maturity as less than one year, and long maturity as more than one year. I then adjust the convenience yield accordingly by taking simple averages over these two maturity categories, and use the bond index for country i at the corresponding maturities as a control for valuation effects.

Since the theoretical model does not feature a maturity dimension, I do not recover the structural parameters from these estimates. However, I can make some heuristic conclusions on the importance of convenience yields. A rigorous estimation of structural parameters would require specifying a theoretical model of the term structure of US and euro area government bonds, which is left for further research.

Figures A1 and A1 display the comparison in yield sensiivity coefficients β for short and long maturities, respectively. For short-term securities, the picture is very similar to the overall portfolio: banks increase the US Treasury share by 5 percentage points in response to a one percentage point increase in excess returns, while insurances only by 1.14. The coefficients are statistically different from each other at the 10% significance level. On the other hand, the estimated reaction for long-term securities is of 1.12 percentage points and not statistically significant for banks, while it is higher at 2.4 percentage points for insurances.

Assuming that the ratio of risk aversion between banks and ensurances remains the same as estimated in section 4, these results suggest that insurances derive a higher convenience yields from short-term than from long-term US Treasuries, consistent with the "moneyness" mechanism.

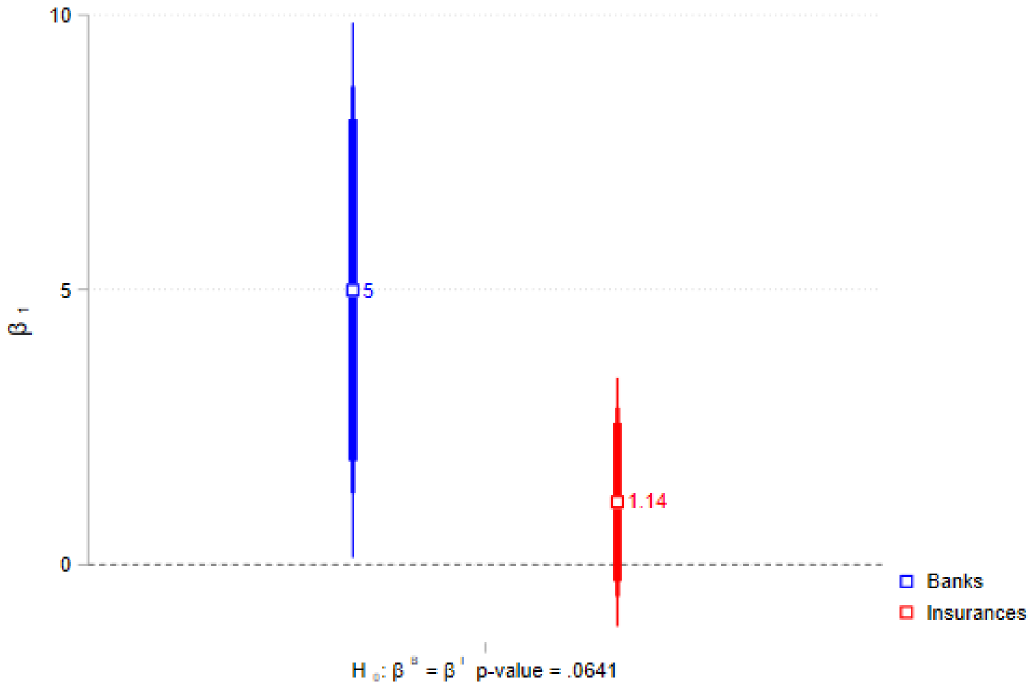
E Robustness checks

E.1 Excluding distressed sovereigns

The baseline specification adopts several precautions to avoid bias stemming from distressed countries in the European sovereign debt crisis of the 2010's. These measures include adjusting convenience yields for CDS rates, controlling for debt/GDP ratios, and country fixed effects.

To judge the effectiveness of this strategy and allay any further concerns of bias from sovereign credit risk, this section repeats the analysis in sections 3 and 4

Figure A1. Comparison of β for banks and insurances - Short maturities

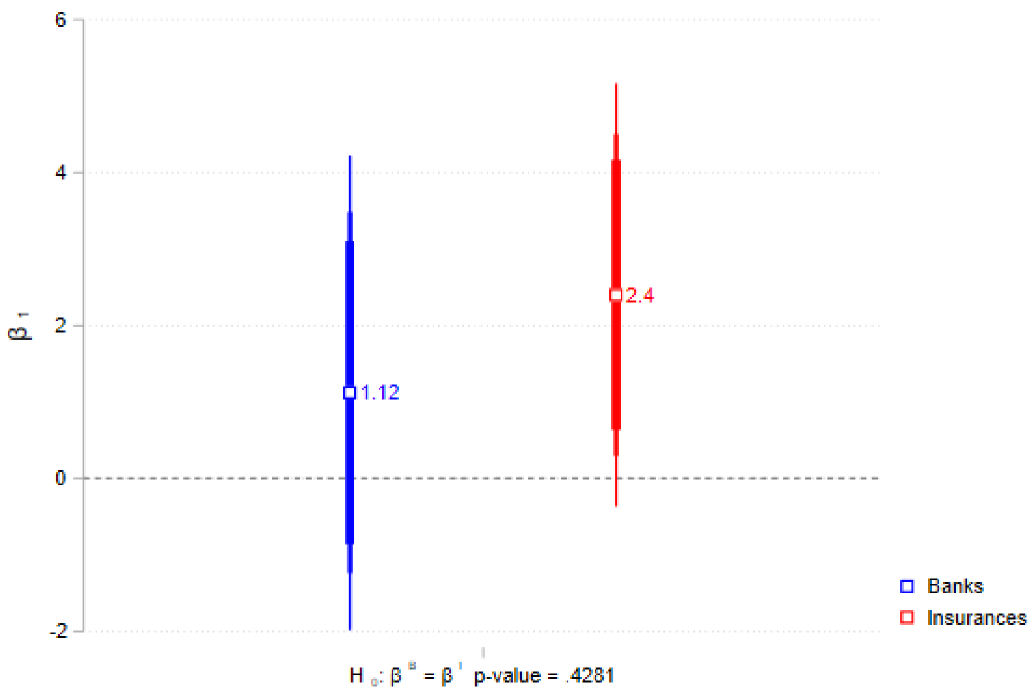


The blue-bordered dot represnets the coefficient β^B, and the red-bordered dot represents the coefficients β^I. Both coefficients are estimated via 2SLS in the model $s^I_{US,i,t} = \alpha^I_i + \alpha^I_t + \beta^I \phi_{i,t} + \delta^{\mathbf{I}'}\mathbf{V_{i,t}} + \eta^{\mathbf{I}'}\mathbf{W_{i,t}} + \varepsilon^I_{i,t}$ via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The bars around the dots represent confidence intervals at the 90%, 95% and 99% levels, in decreasing order of thickness. Portfolio shares, convenience yields and the index of country i government bonds are calculated for maturities shorter than one year. The p-value on the hypothesis $H_0 : \beta^B = \beta^I$ is calculated using the Clogg et al. (1995) method, which assumes that the coefficients are independent.

excluding distressed sovereigns: Cyprus, Ireland, Italy, Portugal and Spain. Recall that Greece is already excluded in the baseline, as its bonds are not eligible under the PSPP. This procedure results in a smaller sample of 201 country-quarter observations.

Figure A3 shows that the coefficients for have a similar magnitude as in the baseline. The coefficient β is almost three times as large for banks (8.74) as for insurances (2.98). However, the estimates of both the individual coefficients and their difference are less precise. The lower statistical significance is presumably due to the loss of observations for large countries like Italy and Spain, which dis-

Figure A2. Comparison of β for banks and insurances - Long maturities

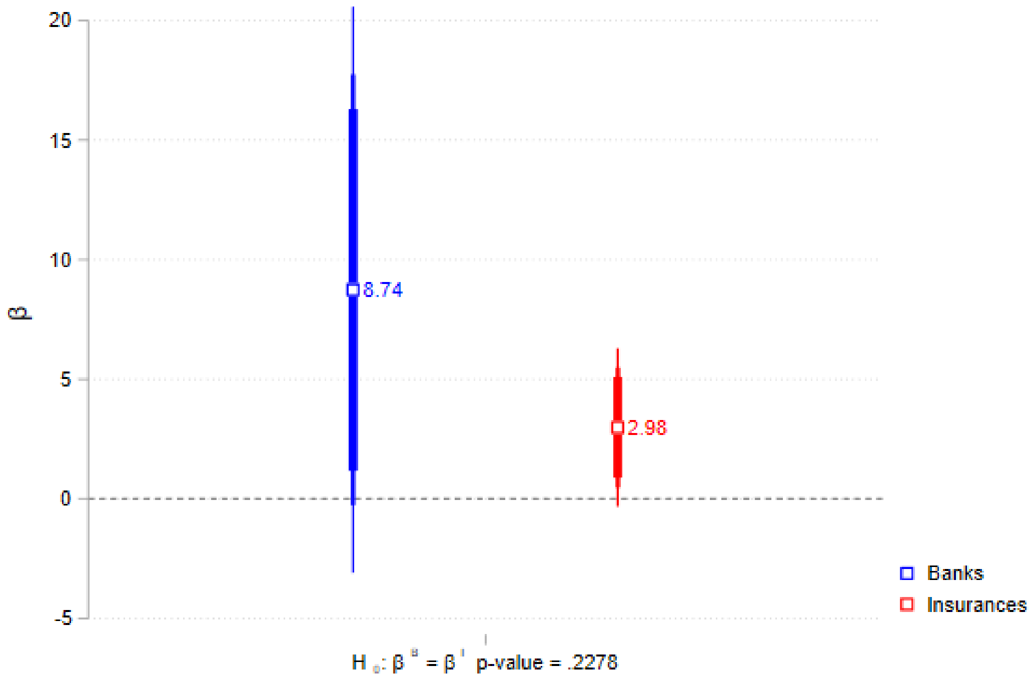


The blue-bordered dot represnets the coefficient β^B, and the red-bordered dot represents the coefficients β^I. Both coefficients are estimated via 2SLS in the model $s^I_{US,i,t} = \alpha^I_i + \alpha^I_t + \beta^I \phi_{i,t} + \delta^{\mathbf{I}'} \mathbf{V_{i,t}} + \eta^{\mathbf{I}'} \mathbf{W_{i,t}} + \varepsilon^I_{i,t}$ via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The bars around the dots represent confidence intervals at the 90%, 95% and 99% levels, in decreasing order of thickness. Portfolio shares, convenience yields and the index of country i government bonds are calculated for maturities longer than one year. The p-value on the hypothesis $H_0 : \beta^B = \beta^I$ is calculated using the Clogg et al. (1995) method, which assumes that the coefficients are independent.

play significant variation in portfolio shares, returns, and PSPP holdings.

Table A3 shows closer estimates in the risk aversion of banks and insurances compared to the baseline. On the other hand, the parameter ψ regulating the weight of non-monetary payoffs in insurers' preferences remains similar. This set of parameters suggests a larger role of convenience yields. This hypothesis is confirmed by figures A4, A5, and A6. The sensitivity of insurance portfolio shares to excess returns would be 80% higher absent convenience yields, which also explain almost 60% of the difference in sensitivity across sectors. Finally, the convenience yield term accounts for more than 10% of the reduction in US Treasury excess returns in response to an increase in eurozone government debt.

Figure A3. Comparison of β for banks and insurances - Excluding distressed sovereigns



The blue-bordered dot represnets the coefficient β^B, and the red-bordered dot represents the coefficients β^I. Both coefficients are estimated via 2SLS in the model $s^I_{US,i,t} = \alpha^I_i + \alpha^I_t + \beta^I \phi_{i,t} + \delta^{\mathbf{I}'} \mathbf{V_{i,t}} + \eta^{\mathbf{I}'} \mathbf{W_{i,t}} + \varepsilon^I_{i,t}$ via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The bars around the dots represent confidence intervals at the 90%, 95% and 99% levels, in decreasing order of thickness. The sample excludes observations from distressed euro area sovereigns, namely Cyprus, Ireland, Italy, Portugal and Spain. The p-value on the hypothesis $H_0 : \beta^B = \beta^I$ is calculated using the Clogg et al. (1995) method, which assumes that the coefficients are independent.

E.2 Excluding the COVID-19 pandemic

The COVID-19 pandemic had a severe impact on financial markets, and on the market for US Treasuries in particular. While money market funds were at the centre of the disturbance in the market for dollar liquidity specifically (Eren et al., 2020), banks and their role in funding sovereigns during crises also received attention (Hardy and Zhu, 2023). More generally, the COVID shock hit banks' profitability and credit rating particularly hard (Aldasoro et al., 2020), raising

Table A3. Structural parameters - Excluding distressed sovereigns

Structural parameter	Mean	95% CI lower bound	95% CI upper bound
τ^B	20.34	2.62	38.13
τ^I	6.24	-3.64	20.44
ψ	0.09	0.0	0.4

Confidence intervals are obtained by drawing 100000 times from the joint asymptotic distribution of parameters in the empirical model, solving for structural parameters for each joint draw, and computing the 5th and 95th percentiles of the simulated distribution.

concerns that their activities during this time might not reflect the ordinary management of their sovereign portfolio. At the same time, life and health insurances were particularly vulnerable to COVID. The associated slump in asset price lowered the value of their assets, while the rapid and substantial increase in mortality and morbidity raised the expected value of their liabilities (Kirti and Shin, 2020).

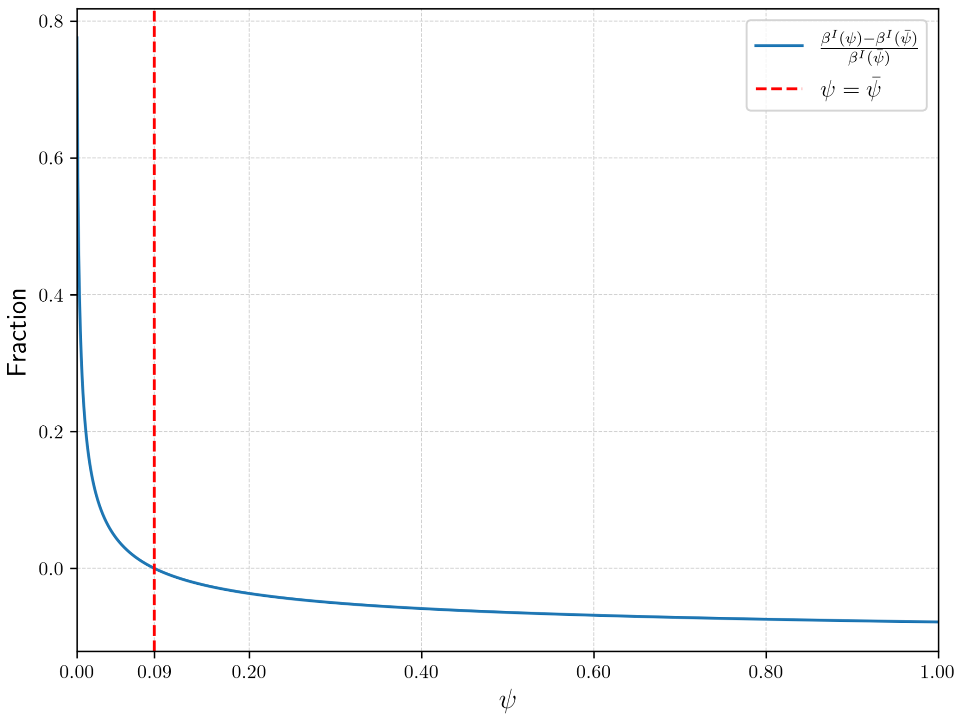
To test the robustness of my results to these unusual circumstances, this section repeats the analysis in sections 3 and 4 excluding all observations after 2020 Q1. This procedure results in a smaller sample of 201 country-quarter observations.

Figure A7 shows that the β coefficients are rather stable, in both size and significance. However, table A4 indicates that the exclusion of COVID substantially raises the risk tolerance of banks relative to insurances. On the other hand, the coefficient ψ on the non-monetary payoff term in insurances' preferences does not change much. Figure A8 shows that, absent convenience yields, the sensitivity of insurance portfolio shares to excess returns would be 50% higher. Figure A9 indicates that slightly less than 50% of the difference in sensitivity across sectors is due to insurers' preference for the non-monetary payoffs of Treasuries. Finally, figure A10 implies that the convenience yield term accounts for not even 1% of the reduction in US Treasury excess returns in response to an increase in eurozone government debt. Overall, these results suggest that the role of convenience yields is smaller if the COVID period is excluded, especially concerning their impact on equilibrium government bond rates.

F Details on the recovery of structural parameters

To solve for the structural parameters, I first proxy $\mathbb{V}[R_{US} - R]$ and $Cov[R_{US} - R, Y^I]$ with their empirical counterparts $\hat{\sigma}_R^2$ and $\hat{\sigma}_{R,Y}$. I also convert the coefficient $-\hat{\pi}$ from the empirical model to account for the minus sign with respect to the

Figure A4. Percentage change in β^I as a function of ψ - Excluding distressed sovereigns



$\beta^I(\psi)$ is calculated using the means of parameters τ^B and τ^I drawn from the simulated distribution, letting ψ vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\psi})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of ψ from the simulated distribution, defined as $\bar{\psi}$.

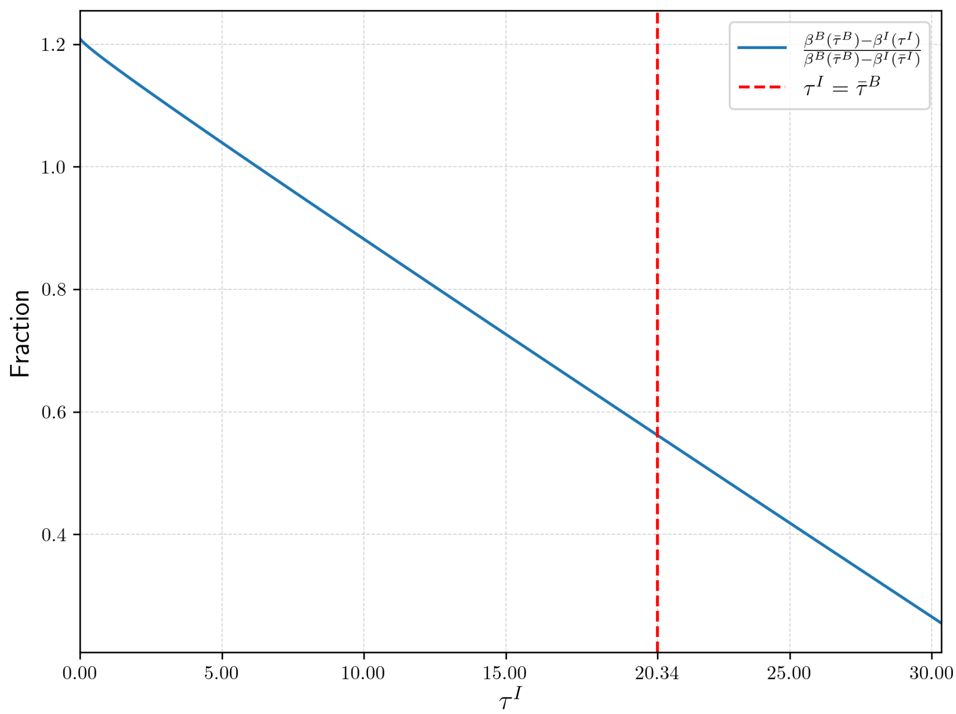
theoretical model, and for the standardisation of PSPP purchases in the empirical model. I then obtain

$$\tilde{\pi} = \hat{\pi}\frac{\bar{B}_{i,t}}{\hat{\sigma}_{PSPP}}$$

where $\bar{B}_{i,t}$ is average outstanding government debt for euro area countries and $\hat{\sigma}_{PSPP}$ is the sample standard deviation of PSPP purchases. Table A5 displays the sample values of the calibrated parameters.

Then, I re-write the system of equations 11, 12 and 13 letting $\tau^B := \frac{1}{\gamma^B}$ and $\tau^I := \frac{1}{\gamma^I}$, to obtain one equation that solves for τ^B directly as a function of ob-

Figure A5. Percentage change in $\beta^B - \beta^I$ as a function of ψ - Excluding distressed sovereigns



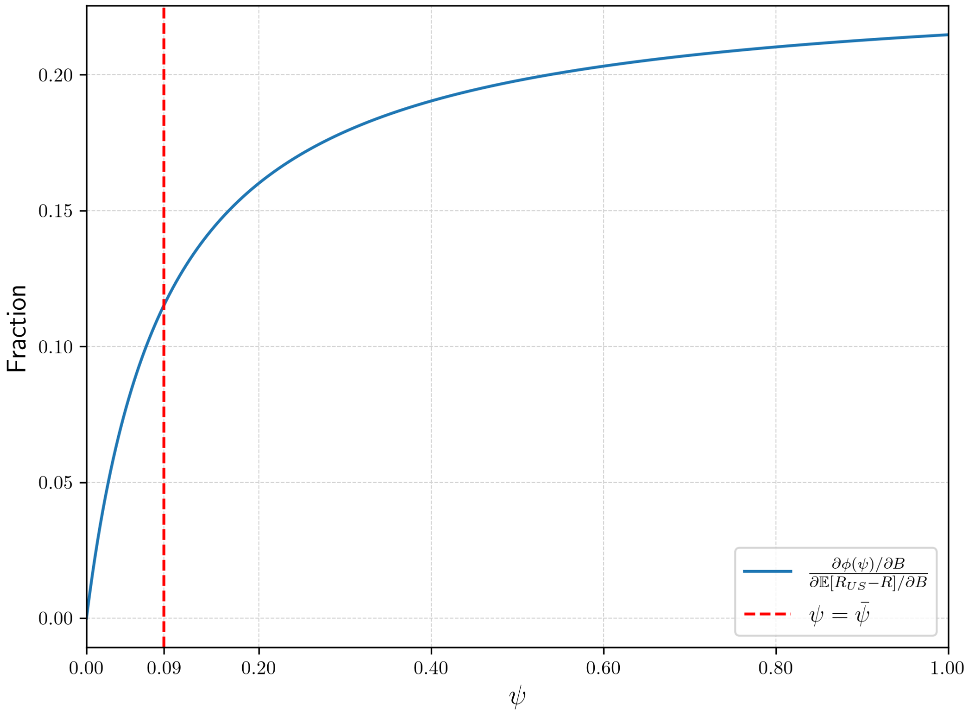
$\beta^I(\tau^I)$ is calculated using the means of parameters τ^B and ψ drawn from the simulated distribution, letting τ^I vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\tau^I})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of τ^I from the simulated distribution, defined as $\bar{\tau^I}$. $\beta^B(\tau^B)$ is calculated using the mean of parameter τ^B drawn from the simulated distribution and the calibrated value for and $\mathbb{V}[R_{US} - R]$ displayed in Table A5.

servables:

$$\tau^B = \hat{\sigma}_R^2 \hat{\beta}^B, \tag{14}$$

and two equations that express ψ as a function of τ^I, τ^B and observables:

Figure A6. Proportion of $\partial\mathbb{E}[R_{US} - R]/\partial B$ explained by the convenience yield component - Excluding distressed sovereigns



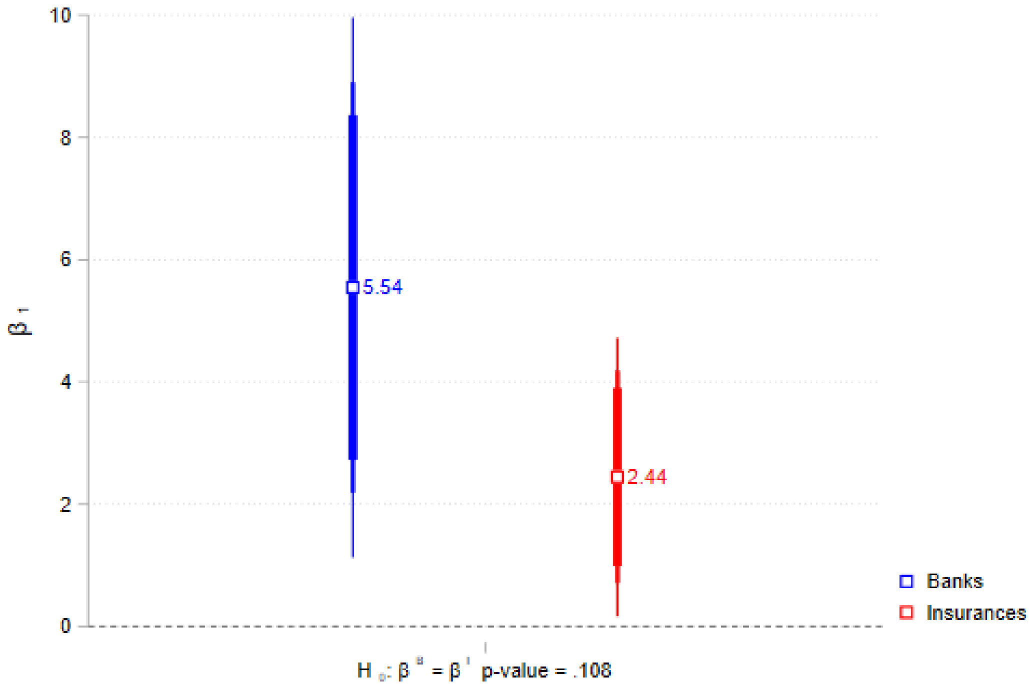
$\beta^I(\tau^I)$ is calculated using the means of parameters τ^B and ψ drawn from the simulated distribution, letting τ^I vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\tau^I})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of τ^I from the simulated distribution, defined as $\bar{\tau^I}$. $\beta^B(\tau^B)$ is calculated using the mean of parameter τ^B drawn from the simulated distribution and the calibrated value for and $\mathbb{V}[R_{US} - R]$ displayed in Table A5.

$$\psi = \frac{(\hat{\sigma}_{R,Y})^2}{4\hat{\sigma}_R^2}\left(\left(\frac{2\hat{\sigma}_R^2}{\tau^I}\hat{\beta}^I - 1\right)^{-2} - 1\right) \tag{15}$$

$$\psi = \frac{0.25\hat{\sigma}_R^2\tilde{\pi}\left(\frac{\tau^B}{\tau^B+\tau^I} - 1\right)^{-1}}{\hat{\sigma}_R^2\tau^I - \tau^B\tilde{\pi}\left(\frac{\tau^B}{\tau^B+\tau^I} - 1\right)^{-1}}. \tag{16}$$

I subtract 15 from 16 and solve the resulting equation for τ numerically using

Figure A7. Comparison of β for banks and insurances - Excluding COVID-19



The blue-bordered dot represnets the coefficient β^B, and the red-bordered dot represents the coefficients β^I. Both coefficients are estimated via 2SLS in the model $s^I_{US,i,t} = \alpha^I_i + \alpha^I_t + \beta^I \phi_{i,t} + \delta^{\mathbf{I}'}\mathbf{V_{i,t}} + \eta^{\mathbf{I}'}\mathbf{W_{i,t}} + \varepsilon^I_{i,t}$ via 2SLS using $PSPP_{i,t}$ as an instrument for $\phi_{i,t}$. The bars around the dots represent confidence intervals at the 90%, 95% and 99% levels, in decreasing order of thickness. The sample excludes observations from 2020 Q2 that are affected by the COVID-19 pandemic. The p-value on the hypothesis $H_0 : \beta^B = \beta^I$ is calculated using the Clogg et al. (1995) method, which assumes that the coefficients are independent.

Powell's method. Finally, I solve for ψ using equation 15.

To simulate the distribution of τ^B, τ^I and ψ, I start by drawing 100000 times from the joint asymptotic normal distribution of $\boldsymbol{\lambda} := (\hat{\beta}^B,\ \hat{\beta}^I\ \hat{\pi})$, assuming independent coefficients. Then, for each joint draw I solve for the structural parameters as outlined above. Then, I calculate the mean, 5$^{\text{th}}$ and 95$^{\text{th}}$ percentile for τ^B, τ^I and ψ over all values of $\boldsymbol{\lambda}$ that admit a solution of equations 14, 15 and 16. I then use the 5$^{\text{th}}$ and 95$^{\text{th}}$ percentiles as bounds for the simulated confidence intervals.

Table A4. Structural parameters - Excluding COVID-19

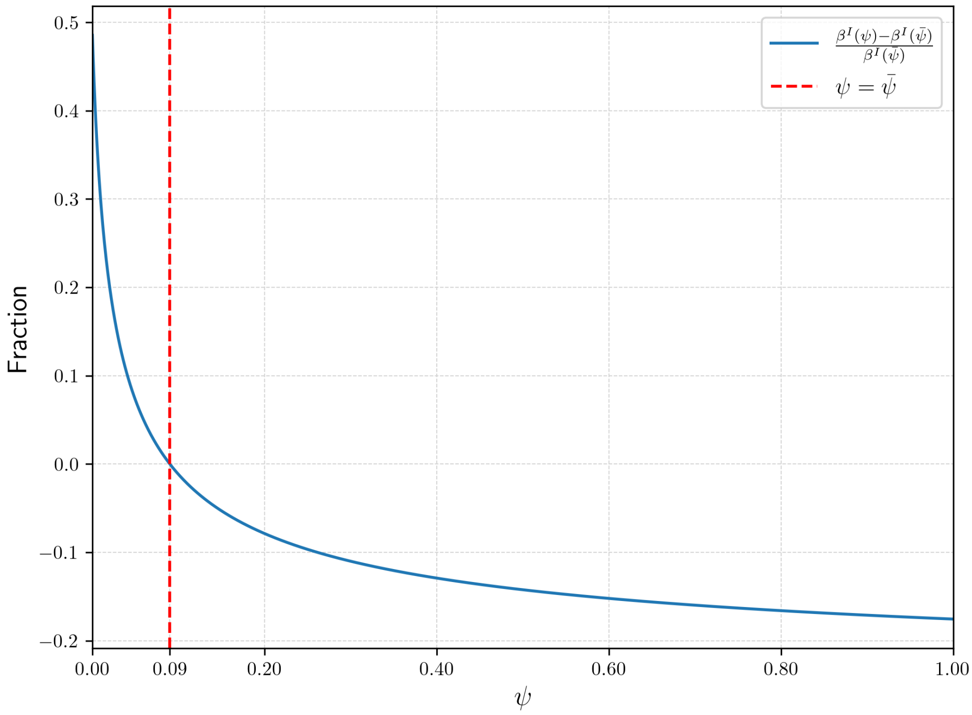
Structural parameter	Mean	95% CI lower bound	95% CI upper bound
τ^B	15.54	7.59	23.52
τ^I	0.63	-16.55	16.55
ψ	0.09	0.0	0.52

Confidence intervals are obtained by drawing 100000 times from the joint asymptotic distribution of parameters in the empirical model, solving for structural parameters for each joint draw, and computing the 5th and 95th percentiles of the simulated distribution.

Table A5. Calibrated parameters

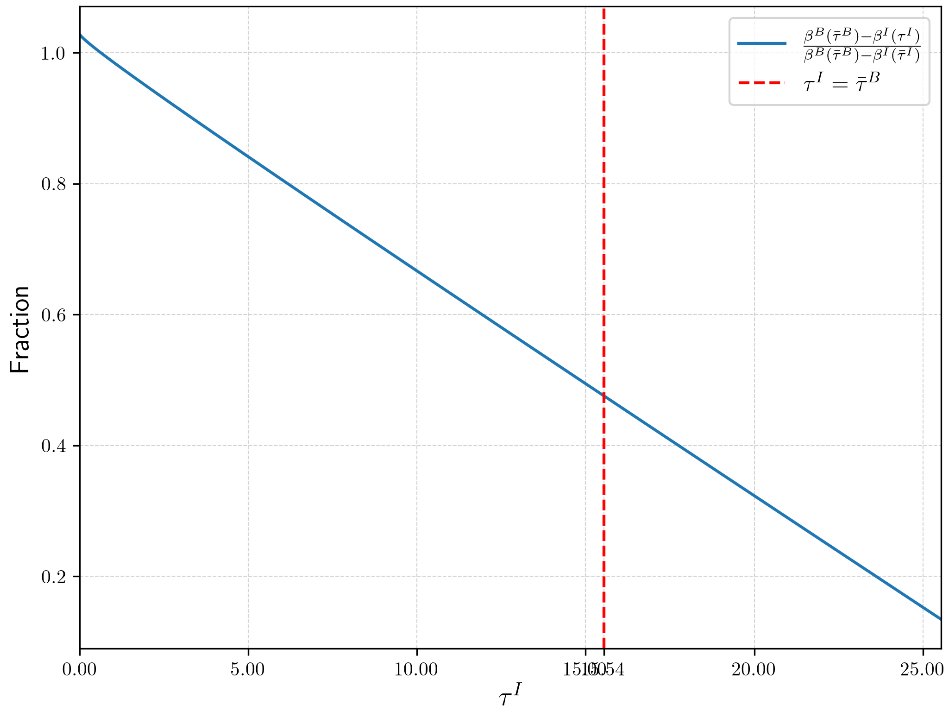
Parameter	Unit of measure	Value
$\hat{\sigma}_R^2$	N.A.	2.22
$\hat{\sigma}_{R,Y}$	N.A.	0.28
$\bar{B}_{i,t}$	€ bn.	527.47
$\hat{\sigma}_{PSPP}$	€ bn.	153.6

Figure A8. Percentage change in β^I as a function of ψ - Excluding COVID-19



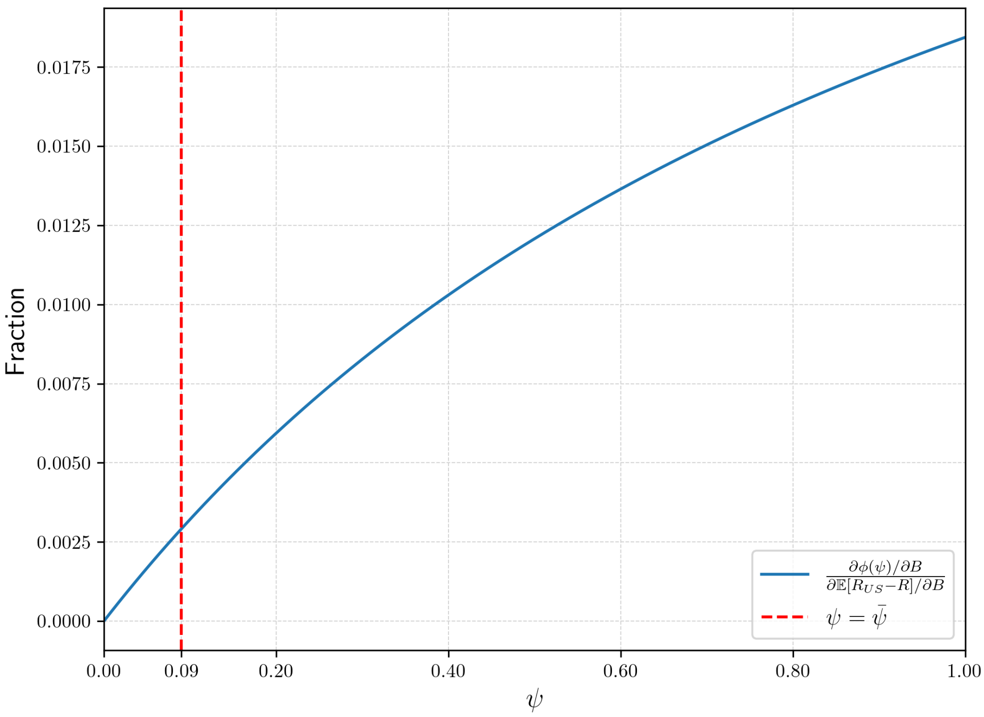
$\beta^I(\psi)$ is calculated using the means of parameters τ^B and τ^I drawn from the simulated distribution, letting ψ vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\psi})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of ψ from the simulated distribution, defined as $\bar{\psi}$.

Figure A9. Percentage change in $\beta^B - \beta^I$ as a function of ψ - Excluding COVID-19



$\beta^I(\tau^I)$ is calculated using the means of parameters τ^B and ψ drawn from the simulated distribution, letting τ^I vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\tau^I})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of τ^I from the simulated distribution, defined as $\bar{\tau^I}$. $\beta^B(\tau^B)$ is calculated using the mean of parameter τ^B drawn from the simulated distribution and the calibrated value for and $\mathbb{V}[R_{US} - R]$ displayed in Table A5.

Figure A10. Proportion of $\partial \mathbb{E}[R_{US} - R]/\partial B$ explained by the convenience yield component - Excluding COVID-19



$\beta^I(\tau^I)$ is calculated using the means of parameters τ^B and ψ drawn from the simulated distribution, letting τ^I vary, and using the calibrated values for $Cov[R_{US} - R, Y^I]$ and $\mathbb{V}[R_{US} - R]$ displayed in Table A5. $\beta^I(\bar{\tau^I})$ is calcualted using the same parameters as $\beta^I(\psi)$, but using the mean of τ^I from the simulated distribution, defined as $\bar{\tau^I}$. $\beta^B(\tau^B)$ is calculated using the mean of parameter τ^B drawn from the simulated distribution and the calibrated value for and $\mathbb{V}[R_{US} - R]$ displayed in Table A5.