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Asset Allocation and Monetary Policy: Evidence from the Eurozone

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Abstract

The eurozone has a single short-term nominal interest rate, but monetary policy conditions measured by real short-term interest rates varied substantially across countries in the period 2003-2010. We use this cross-country variation in the (local) tightness of monetary policy to examine its influence on equity and money market flows. In line with a powerful risk-shifting channel, we find that fund investors in countries with decreased real interest rates shift their portfolio investment out of the money market and into the riskier equity market - causing significant equity price inflation in countries where investment home bias is the strongest.

JEL-Code: G110, G140, G230.

Keywords: monetary policy, asset price inflation, risk seeking, Taylor rule residuals.

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1 Introduction

Following the worst financial crisis (2007–2009) since the Great Depression, a controversial debate has focused on the role of monetary policy for asset price inflation and financial risk-taking in general. Critiques of the U.S. monetary policy have asserted a powerful risk-taking channel whereby excessively low monetary policy rates induce more risky asset allocations by various economic agents (Rajan, 2006; Adrian and Shin, 2010; Borio and Zhu, 2012). Households as well as financial intermediaries might seek higher risk in search of higher yields, and such return chasing may impact leverage and asset prices (Rajan, 2006; Taylor, 2008; Gambacorta, 2009; De Nicolò, Dell'Ariccia, Laeven, and Valencia, 2010). The exceptionally low (and even negative) real short-term interest rate in the current post-crisis environment raises the concern that leverage adjustment is delayed and asset risk allocations are distorted again.

The idea that low real rates for credit may trigger an expansion of leverage accompanied by an asset price boom has a long economic history dating back to Kurt Wicksell (1898).[1] In particular, low real interest rates for riskless investments may entice investors to seek more risky investment positions.[2] Empirically, evaluating the effect of monetary policy on investor behavior faces two types of endogeneity issue. First, the nominal rate setting by a monetary policy authority is a function of business cycle conditions. Such an endogenous nominal rate setting process makes it difficult to determine whether investors react to the nominal rate change itself or to the business cycle condition. The second endogeneity concern is the inflation component of the real rate. Given that local (or national) business cycles can exert influence on local inflation (and hence the local real short rate), any evidence of an association between investor equity flows and the real short rate may be a manifestation of both driven by local business cycles.

This paper seeks to address both endogeneity challenges. To deal with the endogenous

[1] Hellwig (2011) suggests that such Wicksellian dynamics represent a salient feature of southern Europe's recent boom and bust cycle.

[2] The risk-taking channel may operate through (i) increased opportunity costs to investment in low returns assets, (ii) lower investor risk aversion in periods of low real rates, or (iii) less stringent funding conditions for leveraged investments. We do not seek to distinguish these different components.

nominal rate setting, we focus on the eurozone. In particular, we use the monetary policy process in the eurozone with its different national real short-term interest rates to identify how geographic variation in monetary policy conditions affects investors' asset allocations to equity and money market funds.[3] In a currency union, the central bank is constrained to set only one single short-term nominal interest rate for the entire currency area. Therefore, the endogeneity concern is greatly mitigated in our study because we focus on deviations of local monetary policy conditions from eurozone averages (namely, deviations of national real short-term interest rates from eurozone averages). This allows us to explore investors' investment allocations as a reaction to "unintended geographical monetary policy variations." For example, the European Central Bank (ECB) is unlikely to adjust its short-term nominal interest rate just because Spain experiences a higher inflation rate relative to the eurozone average, implying that the nominal rate setting is no longer a function of the local business cycle as it would be under the rate setting by an autonomous Spanish central bank. In particular, the difference between the Spanish real interest rate and the eurozone average is (by construction) orthogonal to the ECB nominal rate setting process.[4]

The second endogeneity issue concerns the local inflationary component of the real rate. Even though the nominal rates are set in Frankfurt, based on euro area aggregates, the Spanish inflation rate itself is affected (or even driven) by the Spanish business cycle. Therefore, any correlation between Spanish fund investors' risk-shifting into equity and a lower Spanish real rate may be a result of changes in aggregate demand (thereafter, the income channel) and/or higher expected local firm cash flows (thereafter, the cash flow channel) in Spain affecting both investor asset allocation and the inflation rate rather than a result of investor risk-shifting

[3] A well-documented strong investor bias toward nationally distributed investment funds (see, e.g., the survey paper by Sercu and Vanpee, 2007) allows us to link local relative monetary conditions to fund-level inflows and outflows in the equity and money markets of different eurozone countries.

[4] We also verify that the ECB's nominal rate-setting process does not affect the *future* real short rate SR asymmetrically across countries in a way that depends on their *current* real short rates. Specifically, we regress the local inflation changes ($\Delta INF_{c,t}$) for each country c at quarter t on lagged euro overnight interest rate changes ($\Delta EONIA_{t-k}$), the real short rate ($SR_{c,t-k}$), and $\Delta EONIA_{t-k} \times SR_{c,t-k}$ in the past one to four quarters ($k = 1, 2, 3, 4$), as well as the country fixed effects. We find no evidence that any of the interaction term $\Delta EONIA_{t-k} \times SR_{c,t-k}$ is statistically significant.

in response to the low real rate itself (thereafter, the risk-taking channel). We employ three empirical strategies to distinguish the risk-taking channel from the two alternative channels.

First, we use control variables that proxy for contemporaneous changes in aggregate output, government spending, and return on assets of local firms to explore whether these variables attenuate the correlation between local real rate changes and local equity fund flows. These income and corporate cash flow measures should represent better proxies for the contemporaneous business cycle than the real short rate does because inflation (and thus the real rate) typically features a more sluggish response to business conditions (due to nominal price stickiness). Inclusion of such control variables in the regression should attenuate the estimated fund flow effect of the real rate if the income and cash flow channels matter much for fund flows. Yet, we find no evidence that these control variables have any significant explanatory power for equity fund flows, whereas the real rate change retains its explanatory power.

Second, under nominal price stickiness, we can instrument the real rate change with its lagged values, thereby restricting the direct influence of contemporaneous business cycle conditions on the estimated fund flow effect of the real rate. Third, we disaggregate equity funds into those with a local investment focus and those with a foreign investment focus. The latter consists of funds that invest more than half of their fund assets in foreign stocks. Such fund flows should not be driven by time-varying cash flow expectations related to local business cycles but rather by business cycles in the foreign investment destination. However, we find that equity fund flows with a foreign investment destination react to the local real rate variations as strongly as flows of funds with a purely domestic investment focus. Taken together, the evidence suggests that investor risk-shifting toward more leveraged equity positions does occur in reaction to changes in the local real rate.

Constrained by data availability, our analysis focuses on investor flows into mutual funds during 2003–2010.[5] Such investor fund flows have a dual interpretation as (i) a measure of investor asset substitution and (ii) a measure of revealed investment and risk preferences. In

[5] Data on other types of money flow in the eurozone (such as hedge fund flows or investment flows for other institutional investors) provides relatively low coverage during our sample period. In light of the data quality concern, we focus on mutual fund flows only.

equilibrium, market clearing implies that net purchases of stocks by local fund investors need to be balanced by corresponding net sales by other local investors or foreign stock investors. Because foreign investors have a different consumption basket, they might not be subject to changes in the local inflation rate and real rate and therefore (ceteris paribus) are likely to accommodate asset demand changes from local investors. In this case, local fund equity inflows can crowd out foreign equity investment and increase equity home bias by local investors. Although our sample is constrained to investor flows to mutual funds, we do find evidence that the local equity mutual fund flows for the eight eurozone countries we use in our sample show a significantly negative correlation with the respective net foreign equity flows of the U.S., consistent with the argument that local fund flows trigger an international asset substitution effect, at least as far as U.S. investors are concerned.[6]

The second and more important interpretation of local fund flows is based on an argument of revealed preference change. Fund inflows are akin to market orders in the market microstructure literature because they represent an investment order for a fixed quantity to be executed (or invested) at an uncertain future price. For example, any buy order is the result of either an increase in the investor's asset valuation or a decrease in the expected execution price. Because correlated buy orders (by a large investor group) can be expected to raise the execution price, any aggregate fund investor inflows need to reflect an even greater change in the (private) equity valuation by this investor group.

We undertake our empirical analysis at the aggregate country level because our variable of interest, the real interest rate, varies only at the country level. Aggregation of fund flow data attenuates flow heterogeneity at the fund level and reduces the role of small funds with

[6]We thank Carol C. Bertaut for providing data on the aggregate foreign equity holdings of U.S. investors, which are estimated based on the methodology described in Bertaut and Tryon (2007). Curcuru, et al. (2011) show that U.S. investors' foreign holdings resemble the market portfolio of the respective country. Therefore, we estimate the quarterly U.S. investor flows into each eurozone member country by first subtracting the product of the beginning-of-quarter holdings and 1 plus the stock market return during the quarter from the end-of-quarter holdings and then scaling the value by the holdings at the beginning of the quarter. We then regress the quarterly aggregate flows of local equity funds (which invest mostly in domestic markets) in the eurozone on U.S. investment flows and time fixed effects. The estimated regression coefficient is -0.27 (t-stat=-4.03).

their more idiosyncratic fund flow patterns.[7] Our results show that loose monetary policy conditions measured by the decrease in the real short rate correlate strongly with the cross-sectional differences in equity fund inflows and money market fund outflows. A decrease of 10 basis points in the real short-term interest rate is associated with a 1% incremental equity fund inflow relative to fund assets and a 0.8% incremental outflow from money market funds.

While fund flow evidence out of money market funds and into equity funds captures an increased risk appetite of a broad investor segment, financial stability concerns the asset price impact of such asset reallocation. We therefore estimate the stock price dynamics triggered by differences in monetary policy conditions in the eurozone using our identification of equity flows related to monetary policy. Accommodating local monetary policy conditions may inflate local equity prices through (i) a lower risk-free (real) rate; (ii) a change in the local risk premia if assets are at least partially subject to local asset pricing; and (iii) a price pressure effect caused by increased equity demand if the asset supply is price inelastic in the short run. Our analysis focuses on the latter two channels by defining for each country a benchmark group of the 20% stocks with the lowest fund holdings over the past three years—called the *Low Fund Holding Index* ($LFHI$). Equity fund returns are measured relative to the returns of this benchmark group and therefore capture the differences in price pressure and/or exposure to changing local risk premia between investable stocks in the fund portfolios and the benchmark low-investability $LFHI$ stocks.

The relative equity fund returns in each country do indeed react positively to local portfolio shifts toward equity triggered by changes in local monetary policy conditions. The measured excess return is approximately 2% for a 10-basis-point decrease in the local real interest rate if all countries are weighted equally. If countries are weighted by the local investment share of domestic institutional investors relative to the local stock market capitalization, we find a much stronger excess return effect of roughly 3.7%—suggesting that the excess return is strongest in countries where local institutional investors are important and exhibit a strong home bias.

[7] Nevertheless, we reproduce our results using fund level regressions and confirm that the coefficients obtained are very similar to those of the aggregate fund flow regressions.

We conduct a number of robustness tests. First, we explore the role of household inflation expectations for the risk-taking channel based on the European Commission's Consumer Survey data. As highlighted by Arnold and Lemmen (2008), collective inflation expectations differ from the best statistical forecast of realized inflation and could represent the more relevant explanatory variable if real investment returns are a key determinant of household risk allocations. In line with this interpretation, we find that after the substitution of the realized real rate changes with the expected real rate changes (calibrated to household inflation expectations), the economic and statistical significance of the real rate effect increases for both equity and money market flows. Second, we conduct a subsample analysis on the pre-crisis period of 2003–2007/q2 and a find qualitatively similar result to our full sample period (2003–2010), alleviating the concern that the recent financial crisis might taint our inferences. Third, we replace the average quarterly $EONIA$ rate with the three-month *Euribor* rate as the proxy for the nominal short-term interest rate and obtain almost identical results. Replacing the real short rate with Taylor rule residuals as the proxy of local monetary policy conditions again yields qualitatively similar results. Fourth, we examine whether inflation-hedging motives can explain our findings of the fund flow effect. Domestic equity investment can be a good hedge against inflation if local inflation and local asset prices move in the same direction. Higher local inflation can also induce the depreciation of the domestic currency and therefore increase the nominal value of foreign assets (after the exchange rate conversion), making foreign equity investment a good hedge against the local inflation risk. However, in a currency union, such as the eurozone, foreign stock investment inside the union does not provide a good local inflation hedge due to the fixed exchange rate arrangement. Yet, our result shows that the correlation between fund flows and local real rates is similarly strong for those funds that invest mostly in other eurozone countries. The evidence does not support an inflation-hedging motive but is consistent with the risk-seeking motive.[8]

Monetary policy is likely to encompass other dimensions than just the short-term rate setting process, such as communicating a long-term policy stance and/or influencing long-term

[8] For a more comprehensive study on household risk-hedging behavior, see Massa and Simonov (2006).

inflation expectations. By focusing on the involuntary cross-sectional differences in the real short rates, we certainly miss any indirect transmission channels common to all countries in the currency union. From this perspective, our study provides a lower bound for the asset allocation effect of monetary policy operating specifically through local real short-term interest rates.

In the following section, we survey the related literature. In Section 3, we discuss the data. Evidence on the asset allocation effect of monetary policy is presented in Section 4. The stock price effect of real rate changes is explored in Section 5. We conclude in Section 6, with some remarks on prudential policies and the stability of a currency union.

2 Related Literature and Policy Issues

The role of asset prices for monetary policy is a subject of considerable controversy. A pre-crisis consensus among many U.S. policy makers was that asset price bubbles were either too hard to identify or beyond the control of monetary policy (Bernanke and Gertler, 1999, 2001; Bernanke, 2002; Kohn, 2006, 2008). An opposing camp argued that a central bank should pay attention to asset price inflation and possibly dampen speculative behavior by increasing interest rates (Borio and Lowe, 2002; Cecchetti, et al., 2000). The latter view is predicated on an endogenous risk hypothesis, whereby investors and/or financial intermediaries seek more risk when real interest rates are low. This view has gained much policy support based on the recent crisis experience although direct empirical evidence for it is still scarce.[9] Yet, such evidence matters not only for the future design of monetary policy but also for gauging the extent to which monetary policy should account for the observed asset price inflation. The current study provides direct empirical evidence on this issue in a unique currency union setting.

The literature has explored a number of risk channels through which loose monetary policy can contribute to financial instability. First, recent evidence supports the view that lax monetary policy affects the riskiness of loans granted by banks (Ioannidou, Ongena, and Peydró, 2009; Maddaloni and Peydró, 2011; Altunbas, Gambacorta, and Marquéz-Ibañez, 2014;

[9] See Issing (2009) for an account of the post-crisis changes in the monetary policy debate.

Jiménez, Ongena, Peydró, and Saurina, 2014). Monetary policy might thus contribute to the build-up of credit risk and bank fragility. Second, low real interest rates might push financial intermediaries to expand their balance sheet and increase their financial risk through leverage (Adrian and Shin, 2010). Our paper focuses on yet another group of investors—retail investors. We argue that these investors might seek more risk in their investment portfolios if low-risk investment provides 'insufficient' returns and renders them less risk averse. A related study by Bekaert, Hoerova, and Lo Duca (2013) provides evidence that innovations to the real interest rate positively correlate with future changes in the VIX index. Such a delayed effect of real interest rates on investor risk aversion is consistent with the direct asset reallocation evidence we document in this paper—real interest rate changes trigger investor preference changes toward less fixed income and more equity investments.

Previous monetary policy research has explored the relationship between nominal rate changes and asset prices. Work by Thorbecke (1997), Rigobon and Sack (2004), Bernanke and Kuttner (2005), and Bjørnland and Leitemo (2009) all document that expansionary (contractionary) monetary policy affects stock prices positively (negatively). Unlike the existing literature, our focus on local fund flows and their disaggregation by investment destination provides more direct evidence for a causal role the real rate plays in investor risk-shifting. Our joint estimation of fund flows related to monetary policy and equity returns also provides a more precise inference of the asset price effect of monetary policy.

Our evidence is also consistent with a large finance literature on the asset price effects of portfolio shifts. For example, Goetzmann and Massa (2003) show how daily S&P500 index returns correlate with contemporaneous index fund inflows. Index fund flows triggered by stock index inclusions or exclusions have been shown to have systematic—though mostly transitory—asset price effects (Chen, Noronha, and Singal, 2004). In our analysis, fund flows are not deemed exogenous; instead, they are examined as a function of monetary policy conditions.

Methodologically, our study benefits from recent advances in the analysis of dynamic panels (Roodman, 2006). Equity fund flows feature a pronounced serial correlation; hence, we need to estimate a dynamic panel for which the ordinary least squares (OLS) or least squares dummy variable (LSDV) estimators are known to deliver inconsistent results—particularly if the time dimension of the panel is small. Our inference is, therefore, based on the use of difference generalized method of moments (DGMM) and system generalized method of moments (SGMM) estimators. We are careful to report the exact instruments set and explore robustness to variations in the instrument choice.

3 Data

A strong home bias in the population of fund investors (who tend to invest in funds that are distributed and marketed locally) allows us to associate local investors' risk choices with inflows and outflows of locally distributed funds. Only investment funds managed in Belgium, Ireland, and Luxembourg appear to draw on a pan-European investor community and therefore are excluded. Greece is excluded because of the lack of fund flow data. Our final sample consists of eight eurozone countries: Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain.

Monetary research has typically inferred a country's monetary policy conditions from either the short-term real interest rate or the so-called Taylor rule residuals. As both the real short rate and the Taylor rule residuals yield very similar results, we focus our analysis on the former and present the result of the latter in the robustness section, 4.5. We measure the quarterly local real short-term interest rate SR by the difference between the average quarterly $EONIA$ rate and the local inflation rate. Real short rates are based on realized quarterly inflation, whereas investors' risk allocation decisions could respond to the expected real rate and expected inflation. Thus, we also derive inflation expectations from the European Commission's Consumer Survey data (see Arnold and Lemmen, 2008, for details). We calibrate the survey response of quarterly inflation expectations to the realized inflation and use the average household inflation forecast to derive the expected real short rate, SR $(expected)$, as described in Appendix A. The correlation between the real short rate and the expected real short rate is high, about 0.97, over our sample period.

Table 1 reports the summary statistics. The average SR $(SR$ $(expected))$ is the lowest in

Spain at -0.096% (-0.046%) and highest in Finland at 0.22% (0.18%) over the 32 quarters of our sample period 2003–2010. Figure 1 plots the time series of the real short rates and the expected real short rates in Panels A and B, respectively, and their changes in Panels C and D, respectively. Overall, monetary policy conditions show considerable independent cross-sectional variation in the eurozone. The average difference between the highest and lowest real interest rate across the eight sample countries is approximately 53 basis points. To measure the local policy conditions relative to the eurozone average in our subsequent analysis, we demean both the real short rate SR and the expected real short rate SR (*expected*) by subtracting from them the respective cross-sectional average rate over the eight sample countries.

The role of local institutional investors also differs across the eurozone countries. Bartram, Griffin, and Ng (2014) document that the average float-adjusted ownership of local institutional investors (reported to the FactSet database) in the quintile of firms with the largest market capitalization value varies from 1.1% for Austria to 10.7% for Germany over the period 2000–2009. We use this ownership share to proxy for the share of the local market held by local institutional investors ($LocInstShare$). We expect that the larger this share, the more likely it is that local equity fund inflows will lead to local asset price inflation.

Our fund flow data are from the Lipper fund database. Fund coverage in Lipper is relatively incomplete prior to 2003. For example, it accounts for only 1.2%, 2%, and 3.3% of the entire mutual fund universe in, respectively, Austria, France, and Germany in 2002 but increases substantially to 60.3%, 68.4%, and 95.7% by the end of 2003.[10] Most funds report returns monthly, but some funds report their total net asset values only quarterly, especially in the early part of our sample period. Therefore, we focus our analysis on the quarterly data from the beginning of 2003 to the end of 2010. Figure 2 contrasts the total fund asset holding statistics reported by Lipper and those reported by the EFAMA. It shows that funds in the eight eurozone countries are generally well represented in the Lipper database, with more discernible coverage

[10] The size of mutual fund industries in the eurozone is obtained from the European Fund and Asset Management Association (EFAMA). It is noted that there are some discrepancies in reporting conventions between EFAMA and Lipper. For example, EFAMA includes funds of funds in the reported statistics of some countries (including France and Italy), but Lipper does not.

shortfall in equity funds for France and Spain and in money market funds for Austria, Italy, and the Netherlands. Such incomplete data coverage may attenuate to some extent the power of our identification mechanism for fund flows in these countries.

To get a cleaner measure of local retail investors' asset allocation reaction to monetary policy conditions, for each sample country we include only funds domiciled and marketed exclusively in the local market. Also, we exclude funds that are sold mainly to institutional investors. Table 2 reports summary statistics for the aggregate fund flows.[11] Across the eurozone, investors generally withdrew capital from money market funds during our sample period. Germany and Portugal experienced the largest outflows, with a mean (median) of -4.8% (-4.0%) and -3.4% (-3.3%), respectively, per quarter. By contrast, investors directed capital into equity funds in Austria, Finland, and Portugal. During this period, equity funds registered a (fund-size-weighted) aggregate mean return of 2.7% per quarter.

Construction of the value-weighted *Low Fund Holding Index* ($LFHI$) uses the semiannual portfolio holdings of worldwide funds from the Thompson Reuters International Fund database. The database is described in detail in Hau and Lai (2013). The 20% least held stocks constituting the $LFHI$ index account for a very small percentage of semiannual holdings by mutual funds. Their aggregate fund holdings relative to shares outstanding range from 0.02% in Portugal to 0.15% in Finland; the average across all eight countries is only 0.07%. Figure 3 illustrates the 20% benchmark $LFHI$ stocks and the remaining 80% of stocks by country in a scatter plot of percentage fund holdings and stock size. The figure shows that the benchmark stocks with extremely low fund holdings exist for a wide range of stock size. The pooled mean return (3.9%) for the $LFHI$ index is close to the return (3.4%) for the MSCI country indices, MKT. We provide detailed definitions and data sources for the aforementioned variables in Appendix A.

[11] The total net asset values of money market funds are completely missing for Finland in Q3 2004 and for the Netherlands in Q4 2002. As a result, Finland has two missing observations for the aggregate money market flows, and the Netherlands has one.

4 Asset Allocation Effect of Monetary Policy

4.1 Base Results

In this section, we examine the relation between local monetary policy conditions across eurozone countries and mutual fund flows into locally distributed equity and money market funds. The serial correlation of fund flows requires us to include a lagged dependent variable in the model specification. A single lagged dependent variable proves sufficient to capture the flow dynamics. We also include lagged market returns ($MKT_{c,t-1}$) in the specification because favorable market returns in a country may correlate with more aggregate equity fund inflows. The regression coefficient of particular interest is α_1, which captures the contemporaneous effect of a country's short-term real interest rate changes ($\Delta SR_{c,t}$) on new equity or money market investment. The specification allows for country fixed effects μ_c and purges time fixed effects by removing the cross-sectional mean from each variable in each quarter:

$$FundFlow_{c,t} = \alpha_1 \Delta SR_{c,t} + \alpha_2 FundFlow_{c,t-1} + \alpha_3 MKT_{c,t-1} + \mu_c + \epsilon_{c,t}. \tag{1}$$

Table 3 presents the regression results for equity funds. Column 1 reports the LSDV estimator as a benchmark, which removes country fixed effects from the regression using the dummy variable approach. Even with the inclusion of country dummies, a short sample of 32 time-series observations suggests that the coefficient estimates are likely to be biased, particularly for the lagged dependent variable. Intuitively, the estimated fixed effects might not fully capture country variations in the average fund flows so that the lagged dependent variable still features some correlation with the residuals, biasing α_2 upwards.

An obvious specification concern is the endogeneity of the real interest rate changes ΔSR to the contemporaneous local business cycle, which may simultaneously influence investor fund allocation decisions and the local inflation rate. The role of the contemporaneous local business cycle effect can be reduced by instrumenting ΔSR and $FundFlow$ with their own lagged values.

A regression based on the DGMM estimator allows for unbiased estimates with the lagged dependent variable, as well as for the instrumentation of covariates. Unlike LSDV, DGMM removes country fixed effects from the data through differencing. Again, we purge time fixed effects by removing the cross-sectional mean from each variable in each quarter. Table 3, Columns 2 and 3, report the DGMM regression results using six and nine instruments, respectively. For $\Delta SR_{c,t}$ and $MKT_{c,t-1}$, we use their own lagged values in the past 1–2 quarters as instruments because they do not feature any autocorrelation at higher orders, whereas for $FundFlow$ we include lags 2–3 of the variable as instruments in Column 2 and lags 2–6 in Column 3.

A comparison of the LSDV estimates with the DGMM estimates shows a slightly smaller coefficient α_2 for the latter. The autocorrelation in fund flows is approximately 0.3 based on the DGMM estimates. A bias-corrected version of the LSDV estimator (not reported) also provides estimates very similar to those in Column 1. However, the use of instruments in Columns 2 and 3 yields a much more negative coefficient estimate for the monetary policy variable. A decrease in the real short-term interest rate by 10 basis points predicts a quarterly equity fund inflow of about 1% of fund assets and a permanent inflow of about 1.4% (estimated by $\alpha_1/(1-\alpha_2)$). These flow effects of monetary policy conditions are therefore statistically highly significant and economically large: If we assume that the flow effect is linear in the real rate changes, then a decrease of one-percentage-point in the real rate corresponds to a substantial 14% of permanent equity inflows relative to fund assets. By contrast, the lagged quarterly aggregate stock market returns, $MKT_{c,t-1}$, do not appear to explain equity fund flows.

An alternative estimation procedure involves the SGMM estimator, which uses both the level and difference equations and estimates both simultaneously. Given the moderate autocorrelation of the lagged flow variable, the SGMM procedure is likely to yield only modest efficiency gains over the DGMM procedure. Moreover, such efficiency gains are achieved only if additional orthogonality conditions for country fixed effects are met (Roodman, 2006).[12] In the interest of a robust inference, we focus our discussions on the DGMM estimates, but report the SGMM results nevertheless in Columns 4–5 using the same instruments used for DGMM1

[12]The orthogonality conditions require aggregate country fund flows to be close to the "steady-state," in which deviations from the long-term values should be orthogonal to country fixed effects after controlling for covariates. It is generally difficult to assert whether such conditions are fulfilled.

and DGMM2 in Columns 2–3. The $\Delta SR_{c,t}$ estimates under SGMM are very similar, but at a slightly higher significance level. The Hansen Test does not reject the validity of the (over-) identification conditions in any of the specifications.[13]

Table 4 provides the corresponding results for money market flows. The estimated autocorrelation for money market flows is between 0.32 and 0.37, similar to that for equity fund flows. The point estimates for the flow effect of the real short rate changes in Columns 1–3 are, respectively, 7.7, 8.5, and 7.8 for LSDV, DGMM1, and DGMM2 specifications, suggesting that a decrease in the short-term real interest rate by 10 basis points predicts a quarterly money market outflow of about 0.8%–0.9% of fund assets. This implies a permanent outflow effect of roughly 1.33% ($\approx 0.85\%/(1-0.36)$) of fund assets.

Overall, the results indicate quantitatively strong equity fund inflows whenever the local monetary policy environment is loose relative to the eurozone average. The corresponding results for money market funds are also economically large, albeit with a lower level of statistical significance.

4.2 The Effect of Local Business Cycles

The evidence of a statistically and economically significant correlation between local real short rates and equity fund flows presented in the previous subsection can have different causal interpretations. In line with a risk-taking channel of monetary policy, low real interest rates may push investors into riskier equity fund investments. Alternatively, the underlying macroeconomic shocks may change aggregate income and corporate profitability, which could simultaneously and directly influence both local inflation and local investor fund flows without a causal linkage from the real short rate to fund flows.

Aggregate demand and supply shocks are the ultimate cause of local inflation (and thus changes in the real short rate), but are they also the "proximate" cause of local investors' risk allocation decisions? Positive aggregate demand shocks increase firm profitability, which could attract net local equity fund inflows. By contrast, negative supply shocks typically generate

[13] The power of the Hansen Test is generally low for a large instrument set. We minimize such a problem by choosing a parsimonious set of instruments.

lower output and decrease corporate profitability; consequently, positive equity fund inflows would occur in parallel with higher inflation only if local investors are (on average) contrarian equity investors. An important source of aggregate demand shocks is local fiscal expenditure. Increased fiscal spending can be inflationary, and at the same time households may decide to save more through equity investment in expectation of higher future taxes.

Local investor reaction to variations in firm profitability, local output, or fiscal spending implies that the inclusion of such macroeconomic variables in the flow regressions of Tables 3 and 4 should attenuate the point estimate for the real short rate and produce statistically significant point estimates for these macroeconomic measures if investor portfolio choice reacts contemporaneously to such real shocks. This argument applies particularly under nominal rigidities, which delay the inflationary effect of macroeconomic shocks and therefore make output, profitability, and expenditure measures a better proxy for contemporaneous macroeconomic shocks than the real short rate, which features a more sluggish adjustment.

In Table 3, Column 6, we augment the baseline regression (DGMM1) by the quarterly changes in local firm profitability, measured by the aggregate return on assets (ΔROA) of locally listed domestic stocks, the national GDP growth ($gGDP$), and fiscal spending growth ($gGovSpd$). The result for equity funds shows that none of these three control variables attenuates the regression coefficient for the real short rate. In particular, the three variables, ΔROA, $gGDP$, and $gGovSpd$, are all statistically insignificant, and the point estimate of ΔSR, -9.58 (t-stat $= -3.58$), is quantitatively similar to the estimate of -9.64 (t-stat $= -3.90$) for the baseline DGMM1 regression.

In Table 4, Columns 6–7, we report the augmented regression results for money market funds. With the inclusion of the three additional variables (ΔROA, $gGDP$, and $gGovSpd$), the point estimates of ΔSR become 7.64 and 8.75, still very close to the estimates of the baseline regressions reported in Columns 2 and 4. Only the coefficient for changes on the return of firm assets (ΔROA) becomes statistically significant. Yet, this effect remains economically small compared to the flow effect captured by the real short rate. We conclude that the fund flow effect we document in Tables 3 and 4 is unlikely to be explained by time-varying expectations about firm profitability over local business cycles.

4.3 Equity Flows by Investment Destination

In this subsection, we examine the equity flows into local funds with different investment destinations (i) to further investigate the business cycle hypothesis discussed in the previous subsection and (ii) to explore yet another alternative hypothesis—the investor inflation-hedging hypothesis. We split the equity fund sample into 15,467 funds investing more than half their assets in domestic equity and the remaining 58,300 funds investing mainly in foreign equity. We then calculate again a country's net aggregate equity fund inflow by its investment destination.

Table 3, Column 7 presents the equity flow regression for funds with a foreign investment focus. The real short rate change, ΔSR, shows a statistically significant equity flow effect, with the point estimate of -11.35, which is slightly larger than the estimate for the full sample. Figure 4 illustrates the negative relationship between the predicted (instrumented) component of the real short rate change and the quarterly aggregate equity fund inflows with a domestic (foreign) investment focus in Panel A (B). The negative relationship extends from mostly domestically invested flows to mostly foreign invested equity flows and is even stronger for the latter. The strong effect of the local real rate change on investment flows into foreign equity is difficult to reconcile with a direct pull effect from the local business cycle because improved cash flow expectations in a local boom should primarily trigger flows into funds with a local investment focus but not funds with a foreign investment focus.

We can investigate further the subset of foreign invested funds whose foreign assets are strictly confined to the eurozone.[14] The estimates, reported in Table 3, Column 8, show that the coefficient of the real short rate changes for this subset of funds is similar to that for all foreign invested funds—albeit at a lower level of statistical significance due to a reduced number of fund observations. This evidence suggests that inflation-hedging motives are unlikely to provide a good explanation for the fund flow effect we document in this paper. While local equities (as real claims) can be expected to increase in price under local inflation and therefore

[14] The information on a fund's investment focus is based on data obtained from Lipper as of December 2010.

serve as an inflation-hedging vehicle, this hedging benefit is largely absent for foreign stocks or eurozone stocks. Intra-eurozone investments in particular are undertaken at a nominally fixed exchange rate, making them a poor hedge against local inflation. A hedging motive should therefore imply a much weaker linkage between the real short rate and the equity flows into funds with a foreign (and particularly eurozone) investment focus, an implication inconsistent with the evidence reported in this subsection.

4.4 Inflation Expectation

So far, we have used the realized real rate changes as the explanatory variable of interest. Using lagged realized changes as instruments (in the DGMM and SGMM regressions) means that we effectively use the predictable component of the real rate change as a regressor. Such an approach is appropriate if investors generally learn about the realized inflation with a quarter's delay. In this subsection, we go one step further and estimate the *expected* real rate change more precisely based on the data from European Commissions Consumer Survey. In particular, we calibrate for each country the average household inflation prediction from the national consumer surveys to the quarterly realized inflation process to obtain the expected local inflation rate and then the expected real short rate change [ΔSR *(expected)*].[15]

Tables 5 and 6 replicate Tables 3 and 4 but replace the real short rate with the expected real short rate. Incorporating quarterly household expectations of inflation into the regression generally increases the magnitude for the coefficient of the real rate. For example, the coefficient drops by 30% from -9.58 (reported in Table 3, Column 6) to -12.37 in Table 5, Column 6. This suggests that the expected inflation component captured by the consumer survey data helps to explain the risk shifting into equity. A qualitatively similar result is obtained for the money market flow regressions. The regression coefficient for ΔSR *(expected),* reported in Table 6, increases relative to ΔSR (reported in Table 4) in economic and statistical significance for every specification reported in Columns 1–7. The overall evidence suggests that the local error component in the household inflation forecast adds explanatory power to both local equity and

[15] See Arnold and Lemmen (2008) for a more detailed analysis of eurozone inflation expectations. We construct the (survey data augmented) expected real short rate using the approach described in Appendix A.

money market flows.

4.5 Robustness

We undertake a variety of robustness checks. First, we verify the stability of our results for pre-crisis period covering 2003–2007/q2. The results reported in Table 7, Panel A, show very similar coefficient estimates for the full sample for both equity fund flows (Columns 1–2) and money market fund flows (Columns 3–4), suggesting that our finding is not driven by the crisis period.

The second robustness test uses an alternative money market rate in the fund flow regressions. In Table 7, Panel B, we construct the real short rate change based on the *Euribor* rate (three-month euro interbank offered rate) instead of the average quarterly $EONIA$ rate. The corresponding point estimates for the coefficient of ΔSR (*Euribor*) are almost identical to those for ΔSR reported in Tables 3 and 4.

Third, we construct local Taylor rule residuals as an alternative measure of the local policy conditions, following the approach used by Maddaloni and Peydró (2011). Table 8, Panel C, shows that the results for changes in Taylor rule residuals (ΔTR) are again qualitatively very similar to those for ΔSR. The numerically larger point estimates for the ΔTR coefficient (e.g., -14.09 and 12.52 for ΔTR versus -9.64 and 8.5 for ΔSR based on the DGMM estimates) reflect the fact that the standard deviation of the Taylor rule residual changes is on average 24% smaller than that of the real short rate changes.

5 Stock Price Effects of Real Rate Changes

5.1 Identification Issues

A major policy concern of low short-term interest rates is asset price inflation, which might result from investor risk shifting from low-yielding fixed income to high-risk equity investment, as documented in the previous section.[16] Unlike the riskless rate effect, which should affect

[16] For example, Jotikasthira, Lundblad, and Ramadorai (2012) show that aggregate fund flows relate to sizeable stock price effects.

assets (of similar duration) alike, the risk-shifting hypothesis of monetary policy predicts that stocks subject to (monetary-policy-related) fund inflows should experience a relatively stronger price appreciation than benchmark stocks of low investability. This implies two identification challenges: First, we need to measure fund returns relative to a local benchmark that is not subject to any asset reallocation effect related to monetary policy. Second, we need to isolate equity fund flows induced by monetary policy conditions from all other (non monetary-policy-related) fund flows.

Fund returns by definition proxy for returns of those stocks in which funds already invest heavily and into which they are likely to channel further investment. In particular, any flow-related price pressure should be captured by the average fund return. By contrast, local stocks of low investability should not be subject to the investor asset reallocation effect (or at least in an attenuated manner) but nevertheless capture changes in the riskless rate and other shocks to the local economy.[17] For each country, we construct a *Low Fund Holding Index* ($LFHI$) based on the returns of the 20% stocks with the lowest fund holdings over the previous three-year period. Because fund flows should primarily impact the returns of the flow-sensitive stocks that constitute the investment universe of the local funds, we can construct a (fund-size-weighted) aggregate local fund return index, $FundReturn_{c,t}$, and identify price pressure as its excess return over the benchmark index $LFHI_{c,t}$ of non-investable stocks:

$$FundReturn_{c,t} - LFHI_{c,t} = \gamma FundFlow_{c,t} + \vartheta_{c,t}. \tag{2}$$

The parameter γ captures the average quarterly return elasticity of fund flows, and $\vartheta_{c,t}$ represents the residual return effects unrelated to fund flows in country c.

The second identifying step involves isolating the (predictable) fund flows induced by the

[17] Importantly, this measure allows us to filter out any unobservable country-wide shocks on firm profitability, which can correlate with monetary shocks. The stock price effect of such macro shocks will not affect our measure unless the cash flow impact of such shocks affects the benchmark and nonbenchmark stocks differently. We verify that both the benchmark and nonbenchmark stocks spread across all industries in our sample, so real shocks are likely to produce similar aggregate stock price impact on both stock samples in each country. Furthermore, the concern that benchmark stocks and nonbenchmark stocks may feature different degrees of liquidity (and thus different expected returns) should not matter for our inference as long as such liquidity differences relate to stock characteristics and do not depend on local monetary policy conditions.

cross-sectional variation in eurozone monetary policy conditions from all other fund flows represented by the residual $\kappa_{c,t}$. In the flow decomposition

$$FundFlow_{c,t} = \widehat{FundFlow}_{c,t} + \kappa_{c,t}, \tag{3}$$

we can use the coefficients estimated from the flow regressions to obtain the predicted fund flows that are triggered by changes in short-term real interest rates as follows:

$$\widehat{FundFlow}_{c,t} = \alpha_1 \Delta SR_{c,t} + \alpha_2 \widehat{FundFlow}_{c,t-1} + \mu_{c,t}, \tag{4}$$

where the coefficients α_1 and α_2 correspond to the estimates obtained in Eq. (1). To derive the predicted fund flows strictly from changes in short-term real interest rates, we drop the market returns from the equation. Similarly, we can further relate $\widehat{FundFlow}_{c,t-1}$ to lagged changes of short-term real interest rates. Substitution of Eqs. (3) and (4) into Eq. (2) yields the specification

$$FundReturn_{c,t} - LFHI_{c,t} = \beta_0 + \beta_1 \Delta SR_{c,t} + \beta_2 \Delta SR_{c,t-1} + \nu_j + \varepsilon_{c,t}, \tag{5}$$

with linear constraints $\beta_1 = \gamma\alpha_1$ and $\beta_2 = \gamma\alpha_1\alpha_2$, and lagged terms $\Delta SR_{c,t-k}$ with $k > 1$ ignored. Eq. (5) can be estimated simultaneously with Eq. (4) under the constraint $\beta_2 = \alpha_2\beta_1$. The sum of the constrained coefficients, β_1 and β_2 , directly reveals the cumulative return effect of changes in short-term real interest rates and thus identifies the role of the risk-shifting channel of monetary policy on the equity prices of those stocks with strong fund flows.

5.2 Evidence

Table 8 provides the estimation results for equations (4) and (5). In Columns 1–4, we report regressions in which each country has the same regression weight, 1/8. Because the share of the local capital market held by local institutional investors ($LocInstShare(c)$) varies greatly for our sample, from 1.1% in Austria to 10.7% in Germany, we expect the fund flows from local investors identified in Eq.(5) to have a significantly larger price impact in Germany than in Austria. Therefore, in Columns 5–8, we use $LocInstShare(c)$ as the country weight to better capture price pressure impact, and we expect the estimated coefficients β_1 and β_2 in Eq. (5) to increase in this case. We estimate the system of equations both for the real short rate changes ΔSR and for the expected short rate changes ΔSR *(expected)*. Specifications 1, 3, 5 and 7 feature no fixed effects for the second equation, whereas country fixed effects are added in specifications 2, 4, 6 and 8.

Estimation of the first equation is undertaken in first differences similar to the DGMM estimates reported in Tables 3 and 5, Column 2. Overall, the corresponding coefficient for changes in real short rates, ΔSR or ΔSR *(expected)*, ranges from -10 to -15, slightly larger than the previous single-equation estimates of -9.6 (reported in DGMM1 of Table 3) and -12.4 (reported in DGMM1 of Table 5).

In the second equation, we impose the restriction that flows triggered by innovations to the real short rates (ΔSR) have a constant price impact γ over time on contemporaneous fund excess returns. The total excess return effect consists of the sum $\widehat{\beta}_1 + \widehat{\beta}_2$. Under equal country weights in Columns 1–2, the total return effect of ΔSR is approximately $\widehat{\beta}_1 + \widehat{\beta}_2 \approx -20$, implying that a decrease of 10-basis-point in the short-term real interest rate increases the relative valuation of flow-sensitive stocks by roughly 2%. However, the standard errors for the coefficients are large, rendering the t-statistics only marginally significant.

By contrast, the results in Columns 5–6, with the country weights based on local institutional investor share, imply economically and statistically significant price pressure effects, with $\widehat{\beta}_1 + \widehat{\beta}_2 \approx -37$. The point estimates of $\widehat{\beta}_1 + \widehat{\beta}_2$ are even more negative if the real short rate is replaced by the expected real short rate ΔSR *(expected)* in Columns 7–8. The estimate in Column 8 suggests that a decrease of 10-basis-point in the real rate boosts relative fund returns in investable stocks by 4.7%. The results suggest that the equity fund inflows triggered by an accommodating monetary policy have a much larger effect on the stock prices of countries where local institutional investors are important and exhibit large home bias.

As a robustness check, we experiment with variations of the 20% threshold for stock inclusion in the *Low Fund Holding Index* ($LFHI$)—using either a 15% or 25% cut-off. Overall, the quantitative return results of Table 8 become slightly stronger for the 15% threshold and slightly weaker for the more inclusive 25% cut-off, but the results remain qualitatively robust across such modifications.

Overall, the asset price effect of monetary policy appears to be large for eurozone countries. Yet, we concede that the benchmark group of "non-investable" stocks might still be tainted by some (small) simultaneous price pressure. As a result, the total excess return effect we reported is likely to underestimate the overall asset price inflation resulting from an accommodating monetary policy. We also note that monetary policy measures may directly influence the discount factor for both investable and non-investable stocks. This may also contribute to an underestimation of the monetary policy effect on equity prices.

6 Conclusion

The recent financial crisis has put research on financial stability and its determinants back on the center stage. An important and unresolved issue remains the role of monetary policy as a contributing factor to instability, particularly if it is very accommodating. This paper contributes to this research agenda by looking directly at the investor asset allocation process in eight eurozone countries, which feature a tight link between the investment decisions of retail investors and fund flows to equity and money market funds in the respective countries.

First, we find that loose local monetary policy conditions—measured by a decrease in the real short-term interest rate relative to the ECB monetary policy at the currency union level—are associated with a strong investor flow out of money market funds and into equity funds even after controlling for contemporaneous local business cycle shocks. The evidence is equally strong for flows into equity funds with a primarily foreign investment focus, suggesting that changing firm cash flow expectations related to the local business cycle cannot explain the risk shifting into equity investment.

Second, we explore whether the asset reallocation process explained by local monetary policy conditions contributes to equity price inflation. We find that investor asset reallocation toward equity funds triggered by loose local monetary policy conditions generates the greatest stock price inflation in countries where local institutional investors hold a large share of the local stock market. This may not be surprising because asset prices ought to be more exposed to risk shifting in reaction to the local real short rate change in markets where local investors are relatively more important. By contrast, financially open economies are more likely to spread asset price inflation globally.

Overall, we interpret our evidence as support for an economically significant link between monetary policy and investors' asset allocation decisions. Loose monetary policy appears to contribute to investor risk-taking through increased equity investment with local equity price inflation as a consequence. It is often difficult for central banks to identify this monetary policy component of asset price inflation, partly due to high overall stock market volatility. Knowledge about investors' asset allocation decisions can serve as a useful complementary source of information about investor risk choices. A prudential policy framework should therefore monitor asset prices in conjunction with micro-level data on investor risk allocations.

Our study also has implications for issues related to the financial stability of a currency union. While it is clear that a currency union, such as the eurozone, sacrifices local monetary autonomy for the sake of capital mobility and fixed internal exchange rates, it is more controversial if the ensuing variation of local monetary policy conditions inside the currency union also gives rise to financial instability. A recent study by Bordo and James (2014) argues that currency pegs (such as the gold standard or more recently the common currency in the eurozone) augment variations in local monetary policy conditions and thus further financial instability. Our evidence on investor risk-seeking as a function of local monetary policy conditions is consistent with such a view. Importantly, we also find that the relative asset price inflation in national equity markets strongly depends on investor home bias and the extent of international diversification in investor equity holdings. Our result suggests that a high degree of financial integration might be a prerequisite for a stable currency union.

Appendix A. Variable Definitions

Variable	Description	Source
$EONIA$	Quarterly average of the overnight interest rate in the euro area.	Datastream
INF	Quarterly inflation rate.	Datastream
SR	Quarterly short-term real interest rate, calculated as the difference between EOINA and the quarterly inflation rate.	Datastream
SR ($expected$)	The difference between the quarterly EONIA and the quarterly expected inflation rate derived from the European Commission's Consumer Survey data. Each month, consumers in the eurozone countries are asked the following question on future prices (Question 6): "By comparison with the past 12 months, how do you expect consumer prices will develop in the next 12 months? They will (1) increase more rapidly, (2) increase at the same rate, (3) increase at a slower rate, (4) stay about the same, (5) fall, or (6) don't know." Let S_i denotes the proportion of consumers choosing option i. The Balance (BAL) statistic is calculated as: $BAL = S_1 + 0.5 \times S_2 - S_5 - 0.5 \times S_4$. Using the quarterly data from 2003/1–2010/4 for the eight sample countries, we run the following pooled regression: $INF_{c,t+1} = \beta_0 + \beta_1 \times INF_{c,t} + \beta_2 \times BAL_{c,t} + \beta_3 \times INF_{c,t} \times BAL_{c,t} + \epsilon_{c,t}$, where c and t are country and quarter subscripts. We obtain the following estimates: $\beta_0 = 0.001$ [$t = 3.12$]; $\beta_1 = 0.783$ [$t = 18.95$]; $\beta_2 = 0.003$ [$t = 2.54$]; $\beta_3 = 0.029$ [$t = 0.15$]. The total number of observations is 256, and the adjusted R-squared is 0.745. The expected inflation for quarter $t+1$ is then estimated by the fitted value of the regression: $0.001 + 0.783 \times INF_{c,t} + 0.003 \times BAL_{c,t} + 0.029 \times INF_{c,t} \times BAL_{c,t}$.	Datastream and Eurostat

Appendix A continued.

Variable	Description	Source
MKT	Quarterly return on the MSCI country market index.	Datastream
$FundReturn$	Aggregate quarterly value-weighted net fund return; each quarter we calculate the average return of all equity (or money market) funds in a country, with individual fund returns weighted by each fund's beginning-of-period fund TNA.	Lipper
TNA	Total net asset value of a fund.	Lipper
Aggregate $FundFlow$	Aggregate equity (or money market) fund flow for a country; it is estimated by the aggregate net dollar flow of all equity (or money market) funds in a country scaled by these funds' aggregate beginning-of-period TNA. A fund's net dollar flow is estimated by the difference between the end-of-period TNA and the product of the beginning-of-period TNA and 1 plus the current fund return.	Lipper
$gGDP$	Quarterly growth of real GDP.	Datastream
ΔROA	Change in return on assets (ROA) at the country level. $ROA(t)$ is measured by the ratio of the aggregate operating income before depreciation over quarter t to aggregate book assets at the end of the quarter. For any two consecutive quarters, we calculate $ROA(t)$ and $ROA(t-1)$ for the same set of firms and then compute ΔROA as $ROA(t) - ROA(t-1)$.	Compustat Global
gGovSpd	Quarterly growth rate of real government expenditure.	Eurostat and Datastream
Euribor	Quarterly real euro interbank offered rate with a maturity of three months (*Euribor*), calculated as the difference between the quarterly nominal *Euribor* rate and inflation rate.	www.euribor-rates.eu

Appendix A continued.

Variable	Description	Source
TR	Residual of a pooled regression of $EONIA$ on the quarterly real GDP growth and inflation rate, with the constraint that the regression coefficients are the same across the eurozone countries: $EONIA_t = \delta_0 + \delta_1 \times gGDP_{c,t} + \delta_2 \times INF_{c,t} + TR_{c,t}$, where c and t denote country and quarter subscripts. Using the data from 2003/1–2010/4 for the eight sample countries, we obtain the following estimates: $\delta_0 = 0.003$ $[t = 8.48]$, $\delta_1 = 0.009$ $[t = 0.55]$, and $\delta_2 = 0.658$ $[t = 11.78]$. The total number of observations is 256, and the adjusted R-squared is 0.349.	Datastream
$LFHI$	Quarterly return on the value-weighted index of the 20% of stocks with the lowest average fund holdings over the previous three years. Fund holdings are aggregated across all funds and scaled by a stock's shares outstanding.	Thomson Financial and Datastream
$LocInstShare$	Average (free-float adjusted) local institutional ownership for the quintile of firms with the largest market capitalization value. The ownership calculation is based on the pool of domestic institutions that report their asset holdings to the FactSet database. The average is first taken by year from 2000/q1 to 2009/q1 and then across time. We obtain the data from Table A3 of Bartram, Griffin, and Ng (2014).	Bartram, Griffin, and Ng (2014)

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Table 1: Macroeconomic Variables

Reported are the summary statistics of the quarterly overnight interest rates in the eurozone ($EONIA$) and the quarterly real GDP growth ($gGDP$), inflation rate (INF), aggregate change in return on assets (ΔROA), and growth of real government expenditure ($gGovSpd$) for the sample countries. The summary statistics for the time-series average proportion of the local stock market held by local institutional investors ($LocInstShare$) is also reported. The sample consists of Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 2003/1–2010/4. We also report the actual short-term real interest rate (SR) and expected short-term real interest rate (SR (*expected*)) by country as well as their cross-country averages. The cross-country averages of changes in the actual short-term real interest rate (ΔSR) and expected short-term real interest rate (ΔSR (*expected*)) are also reported. All statistics are expressed in percent. Appendix A provides the variable definitions in detail.

Variable	Obs.	Mean	Median	STD	Min	Max
Macroeconomic Variables ×100						
$EONIA$	32	0.562	0.516	0.300	0.086	1.047
$gGDP$	256	0.310	0.472	0.925	−6.036	2.670
INF	256	0.460	0.453	0.272	−0.367	1.204
ΔROA	256	0.004	0.008	0.809	−4.794	5.987
$gGovSpd$	256	1.694	0.472	15.854	−37.047	45.257
$LocInstShare$	8	5.387	4.100	3.969	1.100	10.700
Short-Term Real Interest Rate (SR) ×100						
Austria	32	0.101	0.118	0.246	−0.399	0.506
Finland	32	0.220	0.308	0.239	−0.500	0.548
France	32	0.140	0.126	0.250	−0.312	0.678
Germany	32	0.182	0.193	0.192	−0.221	0.501
Italy	32	0.053	0.031	0.224	−0.293	0.594
Netherlands	32	0.165	0.145	0.259	−0.274	0.672
Portugal	32	0.049	−0.014	0.268	−0.440	0.468
Spain	32	−0.096	−0.155	0.260	−0.480	0.408
All SR	256	0.102	0.101	0.258	−0.500	0.678
All ΔSR	256	−0.016	−0.008	0.117	−0.411	0.333
Expected Short-Term Real Interest Rate [SR (*expected*)] ×100						
Austria	32	0.075	0.124	0.244	−0.432	0.450
Finland	32	0.180	0.258	0.217	−0.531	0.446
France	32	0.134	0.103	0.247	−0.332	0.615
Germany	32	0.158	0.173	0.186	−0.237	0.454
Italy	32	0.126	0.108	0.214	−0.239	0.566
Netherlands	32	0.151	0.171	0.238	−0.330	0.544
Portugal	32	0.021	−0.052	0.256	−0.531	0.431
Spain	32	−0.046	−0.087	0.229	−0.381	0.395
All SR (*expected*)	256	0.100	0.113	0.238	−0.531	0.615
All ΔSR (*expected*)	256	−0.018	−0.005	0.101	−0.326	0.268

Table 2: Aggregate Fund Flows

Reported are the summary statistics for the net equity and money market fund flows at the aggregate country level for eight eurozone countries (Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain) during the sample period 2003/1–2010/4. The aggregate fund flow is the aggregate net dollar flow for all funds in a country scaled by their aggregate beginning-of-period total net asset value (TNA). A fund's net dollar flow is estimated by the difference between the end-of-period TNA and the product of the beginning-of-period TNA and 1 plus the current fund return. Also reported are the MSCI country market index return (MKT) and the value-weighted index return for the 20% of stocks with the lowest fund holdings measured over the previous three year period ($LFHI$). The last row of the table reports the statistics for the aggregate fund-size-weighted local fund returns ($FundReturn$).

Variable	Obs.	Mean	Median	STD	Min	Max
Aggregate Equity Fund Flows						
Austria	32	0.007	0.007	0.041	−0.089	0.104
Finland	32	0.018	0.014	0.038	−0.051	0.102
France	32	−0.008	−0.008	0.013	−0.036	0.022
Germany	32	−0.015	−0.013	0.019	−0.063	0.020
Italy	32	−0.032	−0.017	0.036	−0.133	0.009
Netherlands	32	−0.005	−0.005	0.015	−0.036	0.048
Portugal	32	0.002	0.002	0.045	−0.079	0.133
Spain	32	−0.012	−0.003	0.066	−0.220	0.084
All *Fund Flow*	256	−0.006	−0.006	0.040	−0.220	0.133
Aggregate Money Market Fund Flows						
Austria	32	0.001	−0.018	0.068	−0.110	0.170
Finland	30	0.019	−0.013	0.129	−0.249	0.419
France	32	−0.005	−0.013	0.040	−0.070	0.117
Germany	32	−0.048	−0.040	0.049	−0.173	0.058
Italy	32	−0.024	−0.026	0.040	−0.109	0.055
Netherlands	31	−0.006	−0.004	0.052	−0.164	0.165
Portugal	32	−0.034	−0.033	0.082	−0.218	0.185
Spain	32	−0.031	−0.022	0.046	−0.145	0.056
All *Fund Flow*	253	−0.016	−0.022	0.071	−0.249	0.419
Equity Return Indices and Fund Returns						
MKT	256	0.034	0.031	0.142	−0.432	0.388
$LFHI$	256	0.039	0.032	0.136	−0.369	0.450
Fund Return	256	0.027	0.020	0.111	−0.259	0.306

Table 3: Aggregate Equity Fund Flows and Real Rate Changes

The quarterly country aggregate net inflows into equity funds domiciled and marketed in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 2003/1–2010/4 are regressed on changes in the local real short rate in each country (ΔSR). To eliminate the need for time fixed effects, all variables are expressed as deviations from their cross-sectional means. Column 1 provides the estimate using the LSDV regression. Columns (2)–(3) and (4), respectively, provide the estimates using the difference generalized method of moments (DGMM) and system generalized method of moments (SGMM). Column (5) uses the same setup as Column (2) but includes three additional regressors, ΔROA, $gGDP$, and $gGovSpd$. Column (6) provides the DGMM estimate for the net aggregate equity flows based on funds that invest more than 50% of their fund assets in domestic stocks. Column (7) focuses on the net aggregate flows received by those local funds that invest more than 50% of their fund assets in foreign stocks; Column (8) further restricts local fund flows to those funds that invest more than 50% of their fund assets in other eurozone countries. The regressors are (i) changes in the short-term real interest rate ΔSR; (ii) fund flows at lag 1 given by $FundFlow(-1)$; (iii) the country stock market return in the previous quarter $MKT(-1)$; (iv) changes in aggregate corporate profitability, proxied by changes in return on assets (ΔROA) at the country-level; (v) GDP growth ($gGDP$); and (vi) growth in real government expenditure ($gGovSpd$). All regressions report robust t-statistics in brackets. Also reported are the number of observations (*Obs.*), adjusted R-square for the LSDV regression ($Adj.R^2$), type and total number of instruments used in each specification, p-values for the tests of the first and second order autocorrelations of the residuals [$AR(1)$ and $AR(2)$], and Hansen test for the overidentification conditions. Appendix A provides the variable definitions in detail.

	All Aggregate Equity Flows						Flows with Specific Investment Focus	
							Foreign	Eurozone
Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)	DGMM3 (6)	DGMM4 (7)	DGMM5 (8)
ΔSR	−4.538	−9.638	−9.651	−9.261	−9.663	−9.581	−11.349	−10.159
	[−2.04]	[−3.90]	[−3.96]	[−4.91]	[−5.21]	[−3.58]	[−3.87]	[−1.73]
$FundFlow(-1)$	0.332	0.298	0.315	0.285	0.348	0.303	0.386	0.203
	[4.19]	[2.55]	[2.75]	[2.52]	[3.37]	[2.45]	[3.46]	[4.12]
$MKT(-1)$	−0.052	−0.043	−0.052	−0.039	−0.053	−0.045	−0.031	0.010
	[−1.62]	[−1.40]	[−1.52]	[−1.23]	[−1.66]	[−1.22]	[−1.03]	[0.23]
ΔROA						−0.117		
						[−0.68]		
$gGDP$						0.205		
						[0.35]		
$gGovSpd$						0.003		
						[0.14]		
Obs.	254	246	246	254	254	246	240	224
$Adj.R^2$	0.157							
Instruments								
ΔSR		Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundFlow$		Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-3	Lags 2-3
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
ΔROA						Lag 0		
$gGDP$						Lag 0		
$gGovSpd$						Lag 0		
Total Number		6	9	9	12	9	6	6
AR(1)		0.012	0.010	0.012	0.010	0.012	0.024	0.060
AR(2)		0.545	0.546	0.560	0.506	0.506	0.511	0.654
Hansen Test		0.197	0.372	0.537	0.728	0.192	0.532	0.356

Table 4: Aggregate Money Market Fund Flows and Real Rate Changes

The quarterly country aggregate net inflows into money market funds domiciled and marketed in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 2003/1–2010/4 are regressed on changes in the local real short rate in each country (ΔSR). To eliminate the need for time fixed effects, all variables are expressed as deviations from their cross-sectional means. Column (1) provides the estimate using the least square dummy variable (LSDV) regression. Columns (2)–(3) and (4)–(5), respectively, provide the estimates using the difference generalized method of moments (DGMM) and system generalized method of moments (SGMM). Columns (6) and (7) use the same setup as Columns (2) and (4) but include three additional regressors, ΔROA, $gGDP$, and $gGovSpd$. The regressors are (i) changes in the short-term real interest rate ΔSR; (ii) fund flows at lag 1 given by $FundFlow(-1)$; (iii) the country stock market return in the previous quarter $MKT(-1)$; (iv) changes in aggregate corporate profitability, proxied by changes in return on assets (ΔROA) at the country-level; (v) GDP growth ($gGDP$); and (vi) growth in real government expenditure ($gGovSpd$). All regressions report robust t-statistics in brackets. Also reported are the number of observations (*Obs.*), adjusted R-square for the LSDV regression ($Adj.R^2$), type and total number of instruments used in each specification, p-values for the tests of the first and second order autocorrelations of the residuals [$AR(1)$ and $AR(2)$], and Hansen test for the overidentification conditions. Appendix A provides the variable definitions in detail.

Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)	DGMM3 (6)	SGMM3 (7)
ΔSR	7.711	8.503	7.766	9.090	8.948	7.639	8.750
	[2.18]	[1.91]	[1.63]	[2.54]	[2.60]	[1.69]	[2.85]
$FundFlow(-1)$	0.363	0.361	0.315	0.372	0.326	0.361	0.368
	[5.06]	[5.19]	[4.91]	[6.53]	[5.89]	[4.87]	[6.26]
$MKT(-1)$	0.065	−0.009	−0.006	0.009	0.011	0.001	0.012
	[0.91]	[−0.10]	[−0.07]	[0.11]	[0.12]	[0.01]	[0.17]
ΔROA						−0.288	−0.577
						[−2.00]	[−3.99]
$gGDP$						−0.524	0.267
						[−0.81]	[0.45]
$gGovSpd$						0.076	0.068
						[1.56]	[1.50]
Obs.	249	240	240	249	249	240	249
$Adj.R^2$	0.228						
Instruments							
ΔSR		Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundFlow$		Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-3
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
ΔROA						Lag 0	Lag 0
$gGDP$						Lag 0	Lag 0
$gGovSpd$						Lag 0	Lag 0
Total Number		6	9	9	12	9	12
AR(1)		0.011	0.009	0.008	0.007	0.014	0.011
AR(2)		0.801	0.870	0.897	0.974	0.801	0.957
Hansen Test		0.360	0.330	0.730	0.579	0.706	0.968

Table 5: Aggregate Equity Fund Flows and Expected Real Rate Changes

Similar to Table 3, we estimate the quarterly country aggregate net inflows into equity funds, where the actual real short rate changes are replaced with the *expected* real short rate changes (ΔSR (*expected*)), based on quarterly expected inflation rates derived from the European Commission's Consumer Survey data. All other variable definitions are the same as those in Table 3.

	All Aggregate Equity Flows						Flows with Specific Investment Focus	
							Foreign	Eurozone
Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)	DGMM3 (6)	DGMM4 (7)	DGMM5 (8)
ΔSR (*expected*)	−5.398	−12.390	−12.275	−10.954	−11.292	−12.369	−13.990	−16.683
	[−2.01]	[−5.09]	[−5.07]	[−5.76]	[−6.60]	[−4.67]	[−4.80]	[−2.30]
$FundFlow(-1)$	0.334	0.310	0.324	0.298	0.351	0.315	0.398	0.212
	[4.21]	[2.68]	[2.82]	[2.64]	[3.38]	[2.57]	[3.60]	[4.00]
$MKT(-1)$	−0.051	−0.039	−0.047	−0.037	−0.050	−0.040	−0.027	0.022
	[−1.58]	[−1.18]	[−1.31]	[−1.15]	[−1.52]	[−1.05]	[−0.87]	[0.51]
ΔROA						−0.124		
						[−0.75]		
$gGDP$						0.199		
						[0.35]		
$gGovSpd$						0.003		
						[0.13]		
Obs.	254	246	246	254	254	246	240	224
Adj.R^2	0.296							
Instruments								
ΔSR (*expected*)		Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundFlow$		Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-3	Lags 2-3
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
ΔROA						Lag 0		
$gGDP$						Lag 0		
$gGovSpd$						Lag 0		
Total Number		6	9	9	12	9	6	6
AR(1)		0.011	0.010	0.011	0.009	0.011	0.020	0.057
AR(2)		0.442	0.445	0.475	0.438	0.407	0.375	0.695
Hansen Test		0.203	0.400	0.568	0.712	0.222	0.556	0.519

Table 6: Aggregate Money Market Fund Flows and Expected Real Rate Changes

Similar to Table 4, we estimate the quarterly country aggregate net inflows into equity funds, where the actual real short rate changes are replaced with the *expected* real short rate changes (ΔSR (*expected*)), based on quarterly expected inflation rates derived from the European Commission's Consumer Survey data. All other variable definitions are the same as those in Table 4.

Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)	DGMM3 (6)	SGMM3 (7)
ΔSR (*expected*)	9.683	13.428	12.596	12.713	12.563	12.271	12.058
	[2.26]	[2.38]	[2.19]	[2.77]	[2.78]	[2.08]	[2.78]
$FundFlow(-1)$	0.364	0.365	0.319	0.373	0.324	0.363	0.366
	[5.07]	[5.06]	[4.98]	[6.66]	[5.86]	[4.75]	[6.23]
$MKT(-1)$	0.062	−0.018	−0.014	0.004	0.005	−0.007	0.007
	[0.87]	[−0.19]	[−0.15]	[0.04]	[0.06]	[−0.08]	[0.09]
ΔROA						−0.271	−0.559
						[−1.55]	[−3.53]
$gGDP$						−0.499	0.302
						[−0.75]	[0.49]
$gGovSpd$						0.075	0.068
						[1.50]	[1.46]
Obs.	249	240	240	249	249	240	249
Adj.R^2	0.228						
Instruments							
ΔSR (*expected*)		Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundFlow$		Lags 2-3	Lags 2-6	Lags 2-3	Lags 2-6	Lags 2-6	Lags 2-6
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
ΔROA						Lag 0	Lag 0
$gGDP$						Lag 0	Lag 0
$gGovSpd$						Lag 0	Lag 0
Total Number		6	9	9	12	9	12
AR(1)		0.011	0.008	0.008	0.007	0.014	0.010
AR(2)		0.808	0.875	0.912	0.993	0.816	0.981
Hansen Test		0.394	0.326	0.747	0.567	0.707	0.965

Table 7: Robustness

We repeat the fund flow regression in Tables 3 and 4, Columns (2) and (4), for a pre-crisis sub-sample from the first quarter of 2003 to the second quarter of 2007 in Panel A, for a quarterly real short rate change based on the *Euribor* instead of the *EONIA* rate in Panel B, and the change in Taylor rule residuals (ΔTR) instead of the change in the real short rate ΔSR) in Panel C. The Taylor residual (for each country) follow from a pooled (constrained) OLS regression of the nominal rate change (EONIA) on local inflation and GDP growth for all countries. Columns (1)–(2) and (3)–(4), respectively, report the regressions for the net aggregate equity flows and money market flows.

Panel A: Sub-Sample Analysis (2003-2007/q2)

	Equity Flows		Money Mkt. Flows	
Dep. Variable: Fund Flow	DGMM (1)	SGMM (2)	DGMM (3)	SGMM (4)
ΔSR	−10.962	−11.143	9.979	14.858
	[−2.85]	[−3.00]	[1.05]	[2.41]
$FundFlow(-1)$	0.067	0.162	0.289	0.318
	[0.76]	[2.54]	[2.26]	[2.73]
$MKT(-1)$	−0.054	−0.061	0.017	0.043
	[−1.32]	[−1.14]	[0.13]	[0.33]
Obs.	134	142	128	137

Panel B: *Euribor* as Alternative Nominal Rate

	Equity Flows		Money Mkt. Flows	
Dep. Variable: Fund Flow	DGMM (1)	SGMM (2)	DGMM (3)	SGMM (4)
ΔSR (*Euribor*)	−9.638	−9.261	8.503	9.090
	[−3.90]	[−4.91]	[1.91]	[2.54]
$FundFlow(-1)$	0.298	0.285	0.361	0.372
	[2.55]	[2.52]	[5.19]	[6.53]
$MKT(-1)$	−0.043	−0.039	−0.009	0.009
	[−1.40]	[−1.23]	[−0.10]	[0.11]
Obs.	246	254	240	249

Panel C: Taylor Residuals as Alternative Policy Proxy

	Equity Flows		Money Mkt. Flows	
Dep. Variable: Fund Flow	DGMM (1)	SGMM (2)	DGMM (3)	SGMM (4)
ΔTR	−14.085	−15.035	12.523	11.895
	[−3.90]	[−4.78]	[1.84]	[1.82]
$FundFlow(-1)$	0.298	0.252	0.363	0.365
	[2.39]	[1.89]	[5.20]	[5.97]
$MKT(-1)$	−0.028	−0.022	−0.011	−0.005
	[−0.76]	[−0.57]	[−0.12]	[−0.05]
Obs.	240	248	235	244

Table 8: Equity Fund Flows and Fund Excess Returns Simultaneously Estimated

The first equation relates equity fund flows ($FundFlow$) to lagged fund flows and the contemporaneous change in the short-term real interest rate (ΔSR) (or alternatively the change in the expected short-term real interest rate, ΔSR (*expected*), based on inflation expectations from consumer surveys) and is estimated (as before) using the DGMM approach. The second equation relates fund excess returns, $FundReturn - LFHI$, given in Eq. (5), to contemporaneous and lagged short-term real interest rate changes with the cross-equation restriction implied by the estimated flow dynamics. The second equation is estimated without differencing, uses the same instrument set as the first equation and includes either no fixed effects or country fixed effects. To eliminate the need for time fixed effects, all variables are expressed as deviations from their cross-sectional means. The equity fund flow aggregates are based on all locally distributed and marketed equity funds in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 2003/1–2010/4. Columns (1)–(4) present results based on equal country weights (1/8). Columns (5)–(8) use country weights given by $LocInstShare(c)$, defined as the proportion of the local stock market held by local institutional investors. Thus, each country has a regression weight of $[LocInstShare(c)/\sum_c LocInstShare(c)]$ each quarter. All regressions report robust t-statistics in brackets. Also reported are the number of observations (*Obs.*), and type and number of instruments for the GMM estimates.

	Country Weights							
	Equal (1/8)				$LocInstShare$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dep. Variable Equation 1: $FundFlow$								
ΔSR	−12.629	−12.616			−9.936	−10.066		
	[−3.33]	[−3.34]			[−3.07]	[−3.08]		
ΔSR (*expected*)			−14.592	−14.592			−13.428	−13.646
			[−3.20]	[−3.23]			[−3.30]	[−3.36]
$FundFlow(-1)$	0.232	0.241	0.268	0.271	0.177	0.213	0.220	0.248
	[2.13]	[2.39]	[2.38]	[2.62]	[1.55]	[2.02]	[1.87]	[2.31]
Dep. Variable Equation 2: $FundReturn - LFHI$								
ΔSR	−16.066	−16.202			−31.268	−29.709		
	[−1.68]	[−1.65]			[−2.74]	[−2.81]		
$\Delta SR(-1)$	−3.512	−3.479			−6.132	−6.918		
	[−1.68]	[−1.65]			[−2.74]	[−2.81]		
ΔSR (*expected*)			−21.244	−21.803			−37.204	−37.044
			[−1.76]	[−1.77]			[−2.69]	[−2.77]
ΔSR (*expected*) (-1)			−5.402	−5.615			−8.861	−9.881
			[−1.76]	[−1.77]			[−2.69]	[−2.77]
Sum of ΔSR Coeffients	−19.58	−19.68	−26.65	−27.42	−37.40	−36.63	−46.07	−46.93
Country Fixed Effects	NO	YES	NO	YES	NO	YES	NO	YES
Obs.	246	246	246	246	246	246	246	246
Instruments (Eq.1 and Eq. 2)								
ΔSR	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundFlow$	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
Total Number	5	5	5	5	5	5	5	5

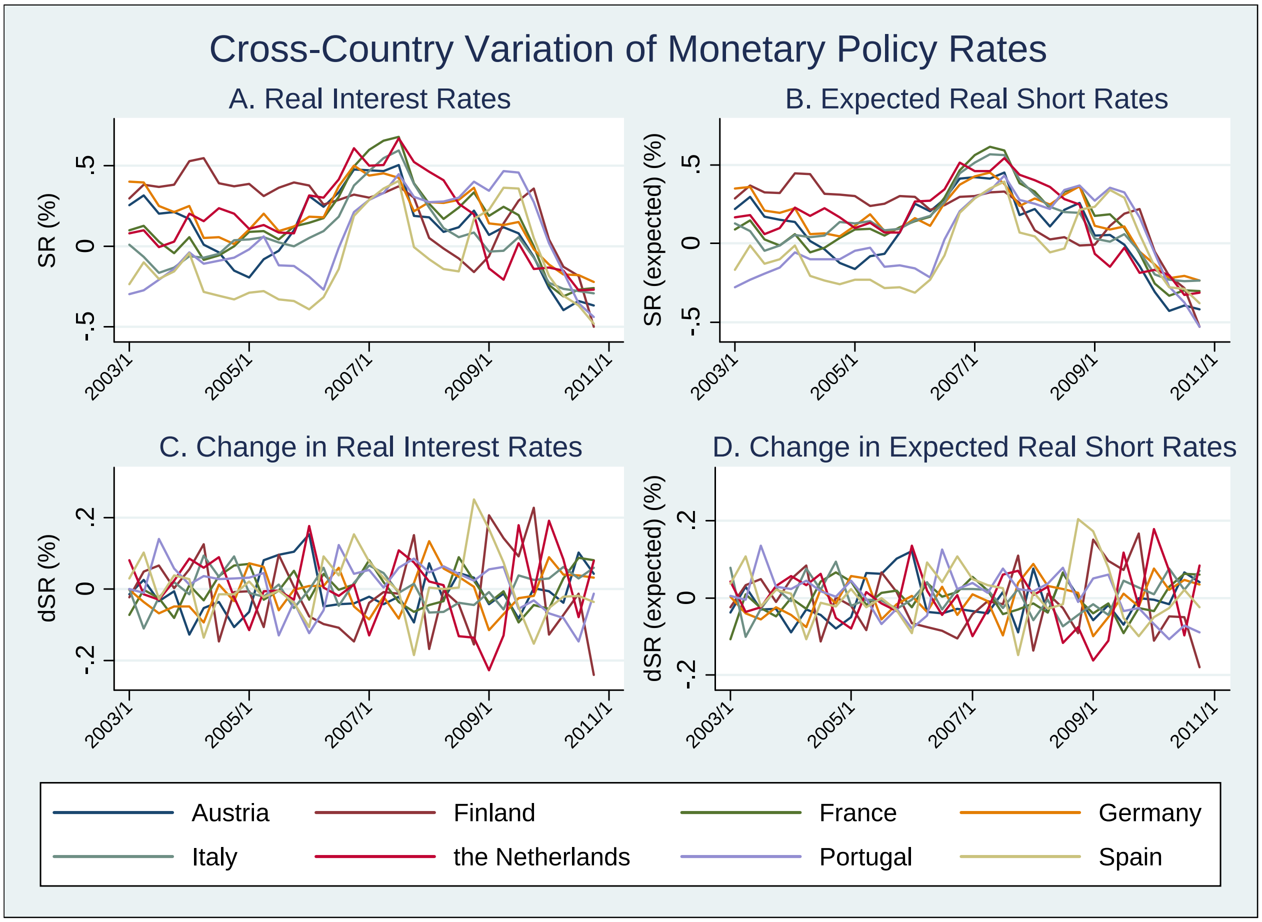

Figure 1: Plotted in Panels A and B are the quarterly real short-term interest rate (SR) and expected real short-term interest rate (SR ($expected$)), respectively, for each of the eight eurozone countries–Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain in the period 2003/1–2010/4. Panels C and D plot the quarterly change of the real short rate (ΔSR) and the quarterly change of the expected real short rate (SR ($expected$)).

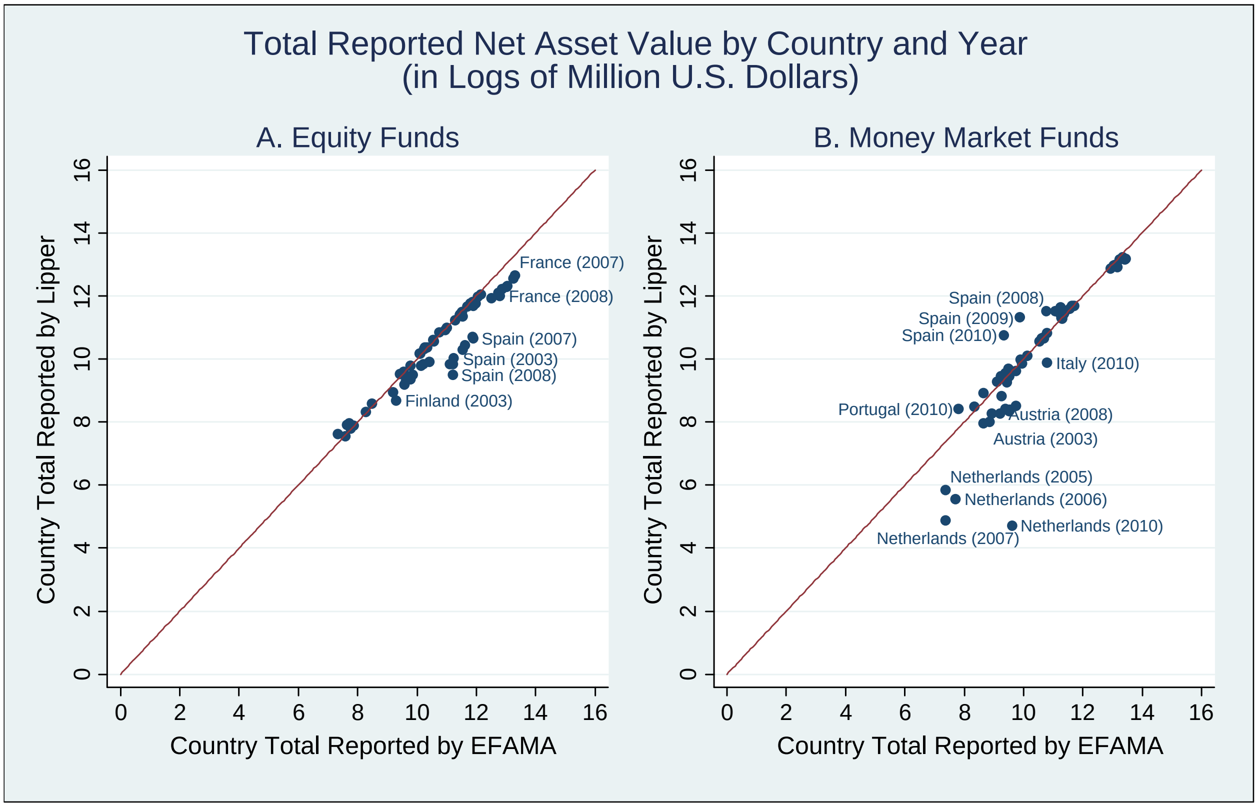

Figure 2: Plotted is the total net asset value (in the natural logarithm of million U.S. dollars) reported by the Lipper fund database on the y-axis against that reported by the European Fund and Asset Management Association (EFAMA) on the x-axis for the eight eurozone countries—Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain—from 2003 to 2010. Panel A plots the equity funds and Panel B the money market funds.

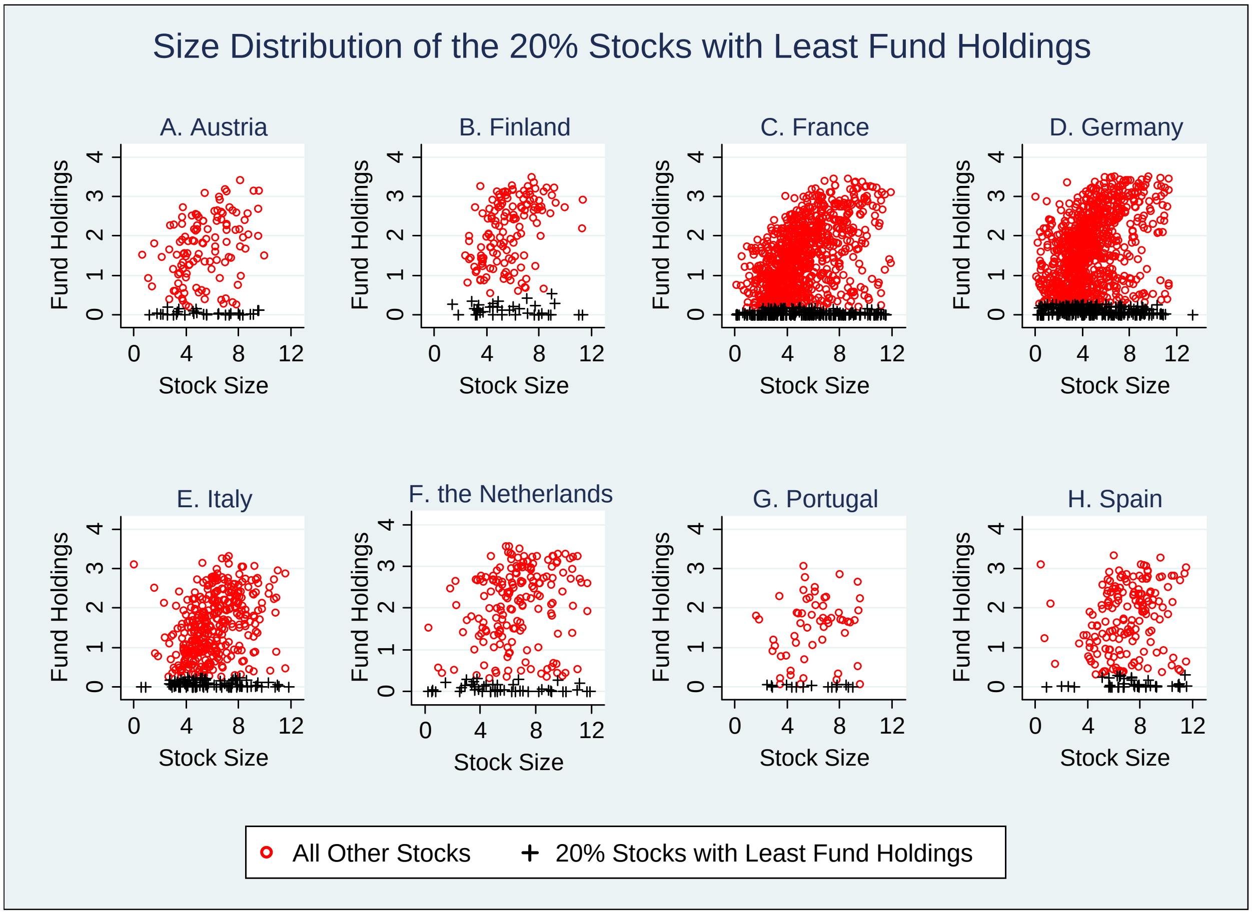

Figure 3: Plotted are the aggregate fund holdings for stocks in eight eurozone countries against the stock size. The 20% of stocks with the lowest fund holdings in each country are marked by black crosses, whereas all other stocks are marked by red circles. We calculate the aggregate fund holdings for each stock as the natural logarithm of the aggregate dollar holdings by all domestic equity funds relative to the stock's market capitalization value at the beginning of the period plus 1, averaged over the sample period 2003/1–2010/4. The x-axis represents the natural logarithm of the market capitalization value of the stock in million U.S. dollars plus 1, averaged over the sample period.

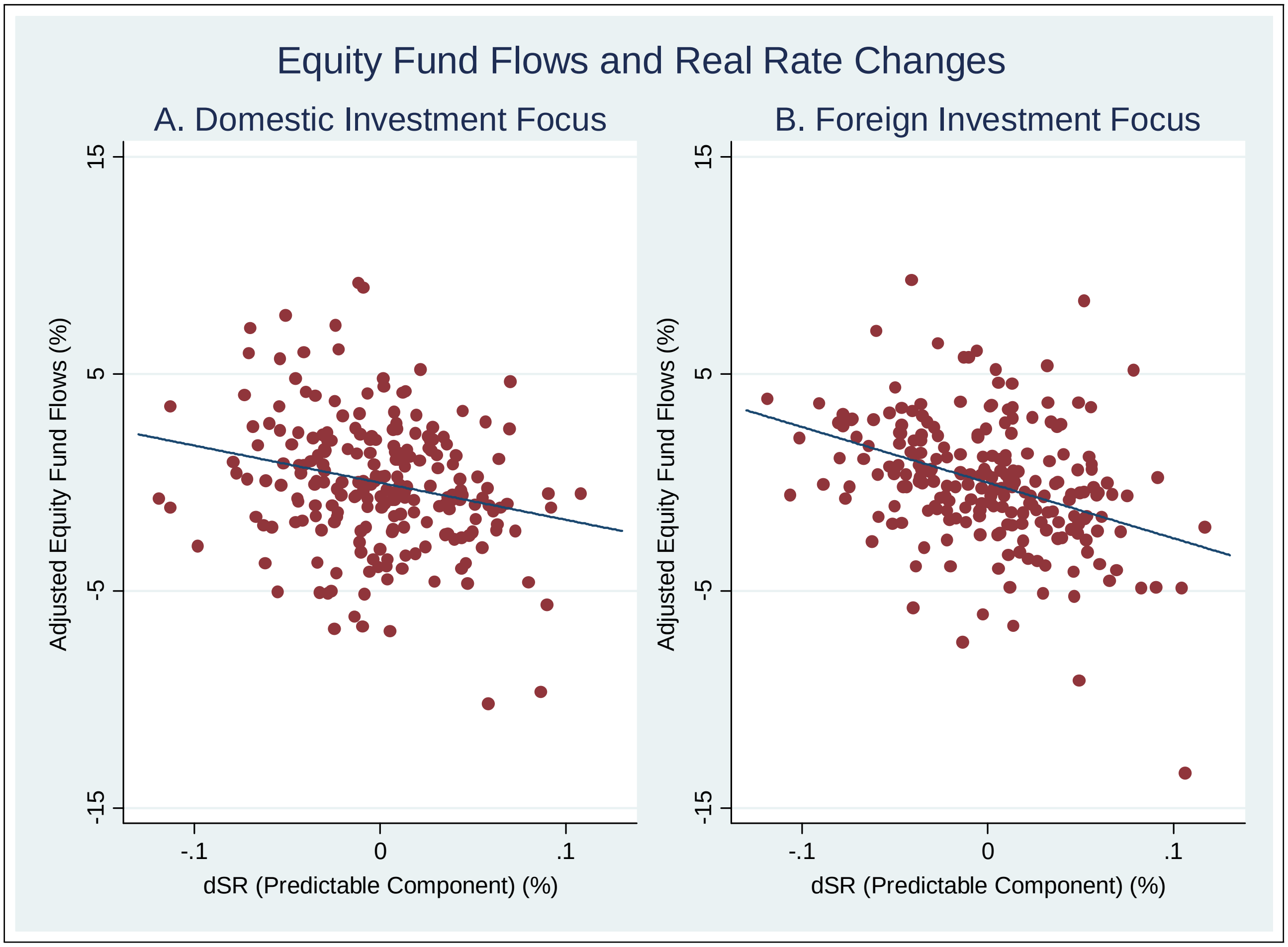

Figure 4: The figure shows the quarterly adjusted equity fund flows (from 2003/1 to 2010/4) for the eight eurozone countries against the quarterly (predicted) change of their respective local real short-term interest rates (ΔSR). Panel A plots the flows for equity funds with a domestic investment focus and Panel B for equity funds with a foreign investment focus. The adjusted equity fund flows denote the difference between the observed equity fund flows and the predictable component of fund flows from lagged fund flow ($FundFlow(-1)$) and lagged market return ($MKT(-1)$). Fund flows are plotted on the y-axis and expressed in percent. On the x-axis, we plot the predictable component of the local real short rate changes (in percent); the predictable component is the one spanned by the instrument set used in Table 3. Panel B is based on the estimates reported in Table 3, Column 7, and Panel A is plotted in a similar manner.