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 either real short-term interest rates or Taylor rule residuals varied substantially across countries in the period from 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. This produces the strongest equity price increase in countries where domestic institutional investors hold a large share of the countries' stock market capitalization.

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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; Borio and Zhu, 2008; Adrian and Shin, 2010). Households as well as financial intermediaries might seek higher risk in search for higher yields, and such return chasing may impact leverage and asset prices (Rajan, 2006; Taylor, 2008; Gambacorta, 2009; De Nicolo, 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.

This paper uses the monetary policy process in the European currency union 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. 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. National equity fund inflows and money market outflows reveal changing investor preferences for asset risk allocation as a function of the local monetary policy conditions. To the best of our knowledge, we are the first to show a strong correlation between local monetary policy conditions in the eurozone and the desire for equity risk exposure by local fund investors. Furthermore, the disaggregate fund flows in each eurozone member country allow us to address the various endogeneity concerns that could not be tackled adequately in the previous macroeconomic research relating monetary policy to stock prices.

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 going back to Kurt Wicksell (1898). Such Wicksellian dynamics appear to be a salient feature of southern Europe's recent boom and bust cycle (Hellwig, 2011). A currency union is likely to feature larger variations in the real rate than any single country—making the eurozone history particularly suitable for the identification of a risk-taking channel of monetary policy.[1] The strong investment home bias and partial integration of European financial markets imply that the asset price effect of household risk shifting is concentrated and identifiable in this setting. Considering that aggregate asset mispricing can still arise under coordination failure in speculative arbitrage (Abreu and Brunnermeier, 2003), asset price booms can translate into financial fragility and pose a threat to the viability of a currency union.

Evaluating the effect of monetary policy on investor behavior faces two types of endogeneity issues. First, the nominal rate setting by the monetary policy authority is a function of the business cycle conditions. Such an endogenous nominal rate setting process makes it difficult to determine whether investors react to a nominal rate change itself or the underlying business cycle condition, which motivates the central bank's nominal rate changes. As nominal rates are set at the currency union level, this endogeneity concern is greatly mitigated in our study. A second endogeneity concern is the inflation component of the real rate, which is affected by the local business cycle. We therefore have to show that household risk allocations react to the local real rate itself, but not to the local business cycle conditions.

To deal with the endogenous nominal rate setting, this paper focuses on the European currency union. In a currency union, the central bank is constrained to set only one single short-term nominal interest rate for the entire currency area. Cross-country differences of either the real short-term interest rate or the Taylor rule residual within the eurozone are orthogonal to the monetary policy process and allow us to explore investors' investment allocations as a reaction to the '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.

The second endogeneity issue concerns the inflationary component of the real rate. Even though the nominal rates are set in Frankfurt based on (exogenous) Euro area aggregates, the Spanish inflation rate itself and hence the Spanish real rate are affected by the Spanish

[1] The average cross-sectional standard deviation of the annual inflation for the eight eurozone countries from 2003 to 2010 is 0.67%, which is more than double the inflation dispersion for the four large U.S. regions over the same period, based on the statistics published by United States Department of Labor.

business cycle conditions. Therefore, any correlation between local fund investors risk shifting into equity and a lower real rate may be a result of changes in aggregate demand (the income channel) and/or higher expected local firm cash flows (the cash flow channel) affecting both investor asset allocation and the inflation rate rather than investor risk shifting in response to the low real rate (the risk-taking channel). We employ three empirical strategies to distinguish the risk-taking channel from the two alternative explanations. First, we use lagged real interest rate changes to instrument the contemporaneous changes. Such instrumentation should restrict the influence of contemporaneous business cycle conditions on the estimated fund flow effect of the real rate changes. Presumably, contemporaneous aggregate demand changes are less correlate with lagged inflation changes (and hence lagged real rates changes) than with the contemporaneous values. Second, we use control variables that proxy for changes in aggregate output, government spending, and return on assets of local firms to explore whether these variables attenuate the correlation between the local real rate changes and the local equity fund flows. Such an attenuation effect should be particularly pronounced under nominal price rigidities because output, government spending, and firm profitability growth should capture asset allocations predicated on the direct income and cash flow channels much better than the (slower) inflation and real rate process. Yet, we find no evidence that these control variables have any explanatory power for equity fund flows, whereas the real rate change retains its explanatory power for fund flows. Third, we examine flows into funds with a foreign investment focus. Such flows should not be driven by the business cycle in the investors' location but rather by the business cycle in the investment destination country. However, we find that flows into these funds react to the local real rate variations as strongly as flows of funds with a purely domestic investment focus do. We conclude that equity fund inflows under low real rates largely reflect increased risk tolerance of local fund investors rather than changes in cash flow expectations.

Previous work on the linkage between monetary policy and equity markets has mainly focused on the effect of asset price inflation (e.g., Bjørnland and Leitemo, 2009). However, any stock price effect alone is uninformative about the risk-taking channel of monetary policy given the endogenous nature of stock prices. By contrast, investor fund flows provide more direct insights into investment preference changes because they can be disaggregated in terms of both investor location and investment destination. Importantly, fund flows have a dual interpretation as (i) a measure of investor asset substitution and (ii) a measure of revealed investment preference (for this investor group). In equilibrium, market clearing implies that net purchases of stocks by local fund investors need to be balanced by the 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.[2]

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 increase the execution price, any aggregate fund investor inflows need to reflect an even larger change in the (private) equity valuation by this investor group.

In terms of eurozone policy measurement, we follow Maddaloni and Peydró (2011) and capture the cross-sectional differences in eurozone monetary policy conditions based on either the local short-term real interest rate (SR) or the country-specific Taylor rule residual (TR). The local real interest rate is defined as the difference between the $EONIA$ rate and the local inflation rate. Alternatively, as explained in Appendix A, we follow Maddaloni and Peydró (2011) to retrieve the Taylor rule residuals from a pooled regression of the common nominal short-term interest rate ($EONIA$) onto the quarterly growth rate for each eurozone country and the corresponding local inflation rate under the constraint of identical coefficients across countries, which embodies the 'average' eurozone Taylor function. We use both the fund level and country level panel regressions to show that loose monetary policy conditions measured by the decrease in either the real short rate or the Taylor rule residual correlate strongly with the cross-sectional differences in equity fund inflows and money market fund outflows. A decrease

[2]For the period from 2003–2010, local equity fund flows for the eight eurozone countries indeed show a significantly negative correlation with the respective net foreign equity flows of U.S. investors into these countries. Local fund flows, therefore, trigger an international asset substitution effect, at least as far as U.S. investors are concerned. See details in Section 3.1.

of ten basis points in the real short-term interest rate (Taylor rule residual) is associated with 1% (1.4%) incremental equity fund inflows relative to fund assets and 0.8% (1.1%) incremental outflows from money market funds. Very similar quantitative results are obtained from panel regressions using either a large cross section of individual fund flows or the aggregation of individual fund flows into country-level flows.

To discard the endogeneity concern that the local business cycle influences not only the local inflation rate, but simultaneously also investor cash flow expectations about locally listed companies, we segment the set of local funds in terms of their investment destination. Our identifying assumption here is that cash flow expectations about foreign firms should not be related to the business cycle conditions in the investor's location. For each country, we identify those funds that invest more than half of their fund assets in foreign stocks. Among these funds, we further examine the subset of funds whose foreign assets are confined strictly in the eurozone. For investor flows of those particular funds, any pull factor emanating from changing cash flow expectations is unlikely to correlate with the inflation rate in the funds' domicile. Yet, our result shows that the correlation between fund flows and local real rates is similarly strong for these subsamples of internationally invested funds—providing support that low local real rates push investors into equity fund investment irrespective of their foreign or local investment focus.

The latter evidence also suggests that inflation hedging motives are unlikely to explain our findings. 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. The equally strong flow evidence for local equity funds with foreign investment focus in the eurozone does not support an inflation hedging motive but is consistent with the risk-seeking motive.[3]

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

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

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 the monetary-policy-related equity flows. Accommodating local monetary policy conditions may inflate local equity prices though (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 through increased equity demand if the asset supply is price inelastic in the short run. Our analysis focuses on the latter two channels by defining in each country a benchmark group of 15% of stocks with the lowest fund flows in the past three years. Equity fund returns are measured relative to the returns of this benchmark group and therefore capture the differences in the price pressure and/or the differences in the exposure to changing local risk premia between the benchmark low-investability stocks and the non-benchmark stocks.

The relative equity fund returns in each country indeed react positively to local portfolio shift toward equity triggered by changes in the local monetary policy conditions. The measured excess return is approximately 1.4% 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 3.4% if the real interest rate is lowered by 10 basis points—suggesting that the excess return is strongest in countries where local institutional investors are important and exhibit a large home bias.

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 inflation expectations. By focusing on the involuntary cross-sectional differences in the real rates and Taylor rule residuals, 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. Giannone, et al. (2011) documents that non-standard monetary measures are employed in some eurozone countries during the recent financial crisis. Yet, our results remain qualitatively unchanged to a more narrowly focused pre-crisis period of 2003–2007/q2, alleviating the concern that such non-standard monetary measures may taint our inferences based on real short rates and Taylor rule residuals.[4] As a second robustness check,

[4] Giannone, et al. (2011) provide a detailed description of the ECB policy during this period. In particular,

we replace the real short rate change with the expected real short rate change; the latter can be inferred from European Commission's Consumer Survey data on inflation expectations. We find qualitatively robust results with respect to the expected real short rate change. Our finding is consistent with the evidence advanced by Arnold and Lemmen (2008) that expected inflation rates in the eurozone countries are strongly determined by past observed inflation.

The following section surveys the related literature. Section 3 discusses identification issues and the data. Evidence on the asset allocation effect of monetary policy is presented in Section 4.1. Section 4.2 addresses the causality issues concerning the relation between fund flows and monetary policy conditions. The stock price effect of investor risk shifting is explored in Section 4.3. Section 4.4 provides robustness tests. Section 5 concludes, 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 its direct empirical evidence is still scarce.[5] 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,

after the Lehman collapse in September 2008, ECB employed some non-standard monetary measures, such as government bond purchases, enhanced credit support, and softening of collateral standards.

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

2009; Jiménez, Ongena, Peydró, and Saurina, 2009; Altunbas, Gambacorta, and Marquéz-Ibañez, 2010; Maddaloni and Peydró, 2011). 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). More leveraged investments by hedge funds might inflate the prices of long positions and expose arbitrage positions to funding risk. Their sudden deleveraging can contribute to considerable asset price volatility and market uncertainty. Third, retail investors might seek more risk in their investment portfolios if low-risk investment provides 'insufficient' returns and renders investors less risk averse. This paper focuses on the last channel and its effect on equity prices.

Bekaert, Hoerova, and Lo Duca (2012) provide evidence that innovations to the real interest rate positively correlate with future changes in the VIX index. They decompose the VIX index into the expected stock volatility and a proxy for the market's risk aversion and show that interest rate changes correlate positively with future variations in the deduced risk aversion. Such a delayed effect of real interest rates on investor risk aversion is consistent with the direct asset reallocation evidence documented in this paper—real interest rate changes trigger investor preference changes toward less fixed income and more equity investments.

Our evidence also relates to a large finance literature that examines 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.

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). Our particular contribution in relation to this strand of literature is threefold: First, examining local policy variation within a currency union (under the single nominal rate constraint) presents a more powerful tool to trace exogenous nominal rate variations, which are much harder to identify convincingly in the traditional single-country macroeconomic analysis. Second, our focus on disaggregate fund flows provides more direct evidence on changing investment preference, and the disaggregation of fund flows by investment destination allows us to disentangle the endogenous cash flow channel from the risk-taking channel of monetary policy. Third, our identification of the asset price effect focuses on the return difference between stocks that receive fund flows due to local real rate changes and those that do not. Joint estimation of these monetary-policy-related flows and equity returns provides a more precise inference on the asset price effect of monetary policy.

Methodologically, our study benefits from recent advances in the analysis of dynamic panels (Roodman, 2006). We measure local investor risk taking based on net equity fund flows of the locally distributed funds. Equity funds feature a pronounced serial correlation; hence, we need to estimate a dynamic panel for which the ordinary least squares (OLS) or least squares dummy variables (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 Empirical Strategies

3.1 Identification Issues

This paper faces three sets of identification challenges, which relate to (i) **the endogeneity of the monetary policy rate setting process**, (ii) the endogeneity of the inflation component of the real rate and the possibility that fund flows may react directly to either aggregate income changes or expectations about firm profitability rather than the real rate itself, and (iii) the quantification of the asset price effect of the fund flows concurring with real rate changes.

To address the endogeneity of monetary policy, we follow the approach used by Maddaloni and Peydró (2011), which exploits the cross-sectional variation of monetary policy conditions in the eurozone. Within the eurozone, there is only one monetary policy and one short-term nominal interest rate across all member countries. Yet, the monetary policy condition differs considerably across nations because of their differences in the inflation rate and GDP growth; euro member countries therefore experience very different real short-term interest rates and Taylor rule residuals. These local deviations in the monetary policy conditions from the eurozone mean are by construction beyond the control of the European Central Bank and hence orthogonal to its policy process. In other words, the institutional constraint of a currency union creates policy-exogenous variations in the monetary policy condition across member countries. For example, even if inflation is two percentage points higher in Spain than the eurozone average, the ECB might not react to such an inflation outlier and leave the nominal rate unchanged, whereas a (hypothetical) autonomous Spanish central bank might increase the nominal rate by two percentage points (resulting in no discernible real rate effects). The real rate differential of two percentage points represents the local real rate variation due to eurozone membership; such a local real rate variation is orthogonal to the ECB nominal rate setting process.

Eliminating the endogenous nominal rate effect by measuring cross-country differences in the real rate is appropriate only if the central bank's nominal rate change does not produce an asymmetric inflation effect across countries systematically. Specifically, the ECB's nominal rate-setting process should not affect the *future* real short rate SR (or Taylor rule residual TR) asymmetrically across countries in a way that depends on their *current* real short rates. We verify this additional orthogonality condition by regressing the local inflation changes ($\Delta INF_{c,t}$) on lagged EONIA 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, indicating that the monetary transmission mechanism does not vary systematically with the 'tightness' of local monetary policy conditions.

Even though the cross section of real rates in the eurozone is exogenous to ECB monetary policy, it is endogenous to the local business cycles. Aggregate demand and supply shocks determine the inflation component of the local real rate. Equity investment decisions by local investors might be affected by the business cycle and changing investment opportunities in the local market. Therefore, local equity fund inflows under low real rates may reflect either investors reacting to higher income (the income channel) and/or higher expected cash flows for local stocks (the cash flow channel) or investors reacting to the low real rate itself (the risk-taking channel). We note that the local business cycle may be larger in a currency union (compared to a counterfactual autonomous central bank that operates its own local monetary policy in the union) because of greater real rate fluctuations. Therefore, any evidence on the linkage between local real rate variations and investor risk-seeking behavior in a currency union is interesting in its own right because of its importance for local financial stability.

We employ three different empirical strategies to distinguish the different channels of the fund flow effect. First, we instrument real short rate changes by their lagged values to filter out the business cycle effect on the contemporaneous real rate changes. Although contemporaneous aggregate income and firm profitability growth may simultaneously increase contemporaneous equity fund inflows and local inflation, equity fund inflows are less likely to correlate with lagged inflation and lagged real rate changes. Second, we directly control for business cycle effects by including in our regressions the local real GDP growth, real government spending growth, and changes in the aggregate return on assets (ROA) of locally listed firms. Such variables should attenuate the role of the real rate in the fund flow regressions if the local business cycle and the related news on firm cash flows are indeed the underlying causes of the observed fund flow-real rate relation. Third, we analyze disaggregate investor flows into equity funds with primarily non-local investment destinations. Such fund flows should not be driven by local business cycles and thus allow us to examine directly whether investor flows react to changes in local real rates.

The interpretation of fund flows in the current paper also deserve some clarification. At the most elementary level, equity flows just measure an investor asset substitution effect because the net supply of stocks is approximately fixed. Given that all European equity markets are highly open to foreign investment, local investors can crowd out foreign investors. We verify this conjecture using quarterly net foreign equity positions of U.S. investors and find that their equity investment flows into the eurozone countries are indeed significantly negatively correlated with the respective countries' local equity fund inflows.[6] Fund flows can also be interpreted as a proxy for changes in the (private) valuation of assets. Investor fund inflows are similar to market buy orders in the microstructure literature because they reveal that this particular investor group has increased its asset valuation beyond the expected execution price.[7] Such an increase in (private) asset valuation can come from either higher cash flow

[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 one 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 (that invest mostly in the domestic markets) in the eurozone on U.S. investment flows and time fixed effects. The estimated regression coefficient is -0.27 (t-stat$=-4.03$).

[7] The expected execution price for market buy (sell) orders is generally above (below) the current midprice

expectations or a lower discount rate due to increased risk tolerance. Therefore, fund flows are of interest not only as a measure of investor asset substitution effect, but also as a measure of the revealed preferences about changing risk tolerance as long as we can discard changing cash flow expectations as the alternative cause of fund flows.

Finally, we seek to identify the linkage between monetary policy conditions and asset price inflation as well as quantify the asset price effect of the enhanced risk seeking by investors. Investor risk shifting in times of low real rates might be only one of the many different factors influencing asset prices. Estimating fund flows and asset prices jointly can help to constrain the analysis and thus provide a more reliable inference on the asset price effect of the fund flows triggered specifically by monetary policy conditions. Generally, three separate channels of monetary policy on asset prices can be distinguished. First, an accommodating monetary policy can set a lower riskless rate (at the short end of the term structure), thus increasing the price of all assets through a lower discount factor. This simple valuation effect may not be a major policy concern and is not the focus of our analysis. Second, changes in monetary policy conditions may change investor risk aversion. An overly accommodating monetary policy may lead to "risk seeking" whereby investors are willing to acquire risky assets at much higher prices. A lower investor risk aversion may rationally explain higher asset prices if the market risk premium (and therefore the discount factor) decreases. Third, any investor asset reallocation to the equity market may generate aggregate mispricing and equity market bubbles. Thus, the asset price inflation may exceed what is predicted by asset pricing models.

Our empirical analysis on asset price effects of monetary policy focuses on the latter two channels by defining for each country, c, a value-weighted *Low Investor Flow Index* ($LIFI_{c,t}$), which aggregates the returns on the 15% of local stocks with the lowest absolute fund inflows and outflows during the previous three years. These particular country return indices focus on the stocks that are least likely to receive additional fund investment. By contrast, fund returns, $FundRetrun_{j,t}$, proxy for the return behavior of the complementary stock universe in which funds invest most. Our analysis of asset price effects is based on the excess return, $FundRetrun_{j,t} - LIFI_{c,t}$, which measures fund returns in excess of the flow-insensitive benchmark return in the respective country. Any change in the riskless rate should equally affect both the fund return and the benchmark portfolio return and is therefore not embedded in this

and particularly so if the market orders are correlated across investors.

excess fund return measure. By contrast, differences in the factor loadings to changing local risk premia as well as differences in the price pressure sensitivity between the benchmark and nonbenchmark stocks should be fully captured by the return difference between the two groups of stocks. Therefore, our excess fund return measure properly identifies the asset price effect of the local equity fund flows triggered by changes in local monetary policy conditions.

Importantly, this measure also allows us to filter out any unobservable country-wide shocks on firm profitability, which can correlate with monetary shocks. For example, local business cycle shocks may create local price inflation and also correlate with future expected firm cash flows. 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.[8] Lastly, 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.

3.2 Data

A strong home bias in the population of fund investors (who tend to invest in those 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 the short-term real interest rate or the so-called Taylor rule residuals. Following Maddaloni and Peydró (2011), we obtain the Taylor rule residuals from a regression of the short-term nominal interest rate on both the GDP growth and inflation rate. A negative (positive) Taylor rule residual at any point in time corresponds to an expansionary (contractionary) monetary policy. For the eurozone, we use a panel regression in which we regress the single short-term nominal

[8] 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.

rate (measured by the EONIA rate) on the GDP growth and inflation rate of all eurozone countries, constraining the regression coefficients to be the same across nations given the single monetary policy. Table 1 reports the summary statistics for macroeconomic variables. The average short-term real interest rate is the lowest in Spain at -0.096% and highest in Finland at 0.22% over the 32 quarters of our sample period from 2003–2010. The alternative measure of monetary policy conditions, Taylor rule residuals, has a high correlation of 0.93 with the short-term real interest rate. Figure 1 plots the real interest rates and Taylor rule residuals in levels in Panels A and B, respectively, and their changes in Panels C and D. 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. The role of local institutional investors also differs across the eurozone countries. Bartram, Griffin, and Ng (2012) report 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 2000-2009 period. 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 is, the more likely local equity fund inflows will lead to local asset price inflation.

Our fund flow data is 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 Austria, France, and Germany, respectively, in 2002 but increases substantially to 60.3%, 68.4%, and 95.7% by the end of 2003.[9] 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 shortfall in equity funds for France and Spain and in money market funds for Austria, Italy, and

[9] 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.

the Netherlands. Such incomplete data coverage may attenuate the power of our identification mechanism for fund flows in these countries to some extent.

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. Our final sample consists of $4,939$ equity funds and $1,441$ money market funds. We estimate a fund's net dollar flow by the difference between its end-of-period total net asset value (TNA) and the product of its beginning-of-period TNA and one plus the current fund return ($FundReturn$). A fund's net quarterly (percentage) flow is calculated as its net dollar flow scaled by the beginning-of-period TNA. We winsorize the 1% highest and lowest outliers of the fund flows in each country-quarter. The aggregate equity (money market) fund flow is the aggregate net dollar flow of all equity (money market) funds in a country scaled by these funds' aggregate beginning-of-period TNA. Table 2 reports fund summary statistics.[10] 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. Across all fund-quarters, the mean (median) flow was 0.8% (-1.1%) for equity funds and -1.5% (-2.7%) for money market funds. The former registered an average quarterly return of 2.3% during this period, compared to 1.1% for the latter.

Construction of the value-weighted $LIFI$ 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 15% least flow-exposed stocks in the $LIFI$ index account for a very small percentage of half-annual fund absolute position changes. Their volume share of total fund trading relative to shares outstanding ranges from 0.02% in Portugal to 0.17% in Finland; the mean volume share over all eight countries is only 0.08%. Figure 3 illustrates the 15% benchmark $LIFI$ stocks and the remaining 85% of stocks by country in a scatter plot of fund flow volume and stock size. The figure shows that the benchmark stocks with extremely low fund flows exist for a wide range of stock size. As reported in Table 2, the pooled mean

[10]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.

(median) return of 3.6% (2.9%) for the $LIFI$ index is about the same as the pooled mean (median) return of 3.4% (3.1%) for the corresponding MSCI country indices (MKT). We provide detailed definitions and data sources for the aforementioned variables in Appendix A.

4 Evidence

4.1 Investment Preferences and Monetary Policy

In this section, we examine the relation between local monetary policy conditions across eurozone countries and mutual fund flows into locally distributed funds. Out of robustness concerns, we present separate evidence on aggregate and disaggregate flows and distinguish in each case between equity and money market flows.

4.1.1 Evidence on Aggregate Fund Flows

First, we report the results for aggregate fund flows, which sum up quarterly individual flows for all funds registered in a country. The serial correlation of fund flows requires us to include a lagged dependent variable in the model specification. For aggregate flow data, 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 reports the regression results for equity funds. Panel A uses short-term real interest rates as the monetary policy variable, whereas Panel B reports identical specifications with Taylor rule residuals as the monetary policy variable. Taylor rule residuals represent estimates with a measurement error, so there may be a concern that our reported regression standard errors are too small for this variable. However, short-term real interest rates do not suffer from this shortcoming.

Table 3, Column 1, reports as a benchmark the LSDV estimator, 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 contemporaneous local business cycle, which may simultaneously influence investor fund allocation decisions and local inflation rate. The role of contemporaneous local business cycle effect can be reduced by instrumenting ΔSR and $FundFlow$ with their own lagged values as both real short rate and fund flows are autocorrelated. Furthermore, equity flows might react not only to real interest rates, but also to the past performance of the local stock market. We therefore include the local stock market return, MKT, in the previous quarter as a control variable. The lagged market return is also instrumented with its own past 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, regardless of whether the short-term real interest rate (Panel A) or the Taylor rule residual (Panel B) is used to proxy for the local monetary policy condition. 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 $\alpha_1/(1-\alpha_2)$). The standard deviation of quarterly changes in Taylor rule residuals is 0.089 (reported in Table 1), which is approximately 24% smaller than the standard deviation of changes in real short rates. Accordingly, we find that a decrease in the Taylor rule residual by 10 basis points corresponds to a quarterly equity inflow of about 1.4% of fund assets and permanent inflows of about 2%. These flow effects of monetary policy are therefore statistically highly significant and economically large: If we assume that the flow effect is linear in the real rate changes, then a one-percentage-point decrease in the real rate corresponds to a substantial 14% of permanent equity inflows. 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 the two equations 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).[11] To be conservative, we focus our discussions on the DGMM estimates but report the SGMM results as a robustness check. Table 3, Columns 4 and 5, report the SGMM results with the same instruments as those for DGMM in Columns 2–3. The $\Delta SR_{c,t}$ estimates under SGMM are very similar to those under DGMM 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.[12]

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, reported in Panel A, are now 7.7, 8.5, and 7.8 for LSDV, DGMM1, and DGMM2, respectively, 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. Using Taylor rule residuals instead of short-term real interest

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

[12] 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.

rates in Panel B again shows that the estimated flow effects are large: A loose monetary policy with the Taylor rule residual lowered by 10 basis points predicts an immediate incremental money market outflow of approximately 1.1% of fund assets and a permanent effect of roughly 1.57% ($\approx 1.1\%/(1-0.3)$).

Overall, the aggregate flow regressions 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. The next section explores whether this finding is robust to the disaggregate analysis at the fund level, which allows for a larger cross section of observations and greater statistical power, as well as for the inclusion of fund-level controls such as fund performance.

4.1.2 Evidence based on Disaggregate Fund Flows

Aggregating individual fund flows to a country-level panel involves a loss of information. Fund-level panels allow for a much larger cross section of $4,939$ equity funds and $1,441$ money market funds instead of the eight eurozone countries. They also allow us to control for fund-level performance, which has been established as an important driver of investor flows (Sirri and Tufano, 1998). The following regression controls for the quarterly lagged fund performance ($FundReturn_{j,t-1}$):

$$\begin{aligned} FundFlow_{j,t} &= \alpha_0 + \alpha_1 \Delta SR_{c,t} + \alpha_2 FundFlow_{j,t-1} + \alpha_3 FundFlow_{j,t-2} + \\ &\quad + \alpha_4 MKT_{c,t-1} + \alpha_5 FundRetrun_{j,t-1} + \mu_j + \epsilon_{j,t}. \end{aligned} \tag{2}$$

Unlike aggregate flows, individual fund flows show significant dependence on the second lag of the dependent variable, which is therefore included in the disaggregate flow specification. Again, we allow for a (fund) fixed effect μ_j and transform both the dependent and independent variables into deviations from their cross-sectional means to remove the impact of time fixed effects.

Because smaller funds may feature higher and noisier flow variability, we reduce their role in the regression by using beginning-of-period fund asset values as regression weights within the group of funds in a country. Value-weighting has the added benefit of making the coefficients in the fund-level analysis more comparable to those in the country-level analysis. We also repeat the analysis using an equal-weighted approach and find similarly strong monetary policy effect on fund flows. We discuss these results in more detail together with other robustness checks in Section 4.4.

Similar to the case for aggregate flows, the lagged dependent variables $FundFlow_{j,t-1}$ and $FundFlow_{j,t-2}$ feature estimation bias if fund fixed effects matter. Therefore, the least squares dummy variables specification in Table 5, Column 1, is biased in spite of the inclusion of fund fixed effects. The difference GMM estimator serves as a useful approach to deal with the estimation bias. The instrument set used in each specification is stated at the bottom of each panel. A comparison of the LSDV result in Panel A, Column 1, to the corresponding DGMM results in Column 2 shows that the former yields an estimated autocorrelation of 0.20 for fund flows, which is substantially lower than the estimate from the aggregate flows (reported in Table 3), suggesting a highly biased LSDV estimate. By contrast, the DGMM specifications yield an estimated autocorrelation of about $0.34 - 0.35$, similar to the estimate using the aggregate flow data. At the disaggregate level, lag 2 of fund flows still enters significantly with a value of 0.13. Lagged market returns, $MKT(-1)$, again have no reliable explanatory power in the DGMM regressions, consistent with the findings from Tables 3 and 4. By contrast, lagged fund returns are a highly significant determinant of equity flows. The more elaborate specification labeled DGMM2 in Table 5 implies that a 1% higher quarterly fund return in the previous quarter correlates with a short-run equity inflow of about 0.2% of asset values.

Of particular interest is the coefficient for change in the real short rate, ΔSR. The fund-level regressions for DGMM in Table 5 yield almost the same equity flow elasticity of about -10 as that in the country-level regression reported in Table 3, but the standard error is now considerably lower. Hence, the relation between loose monetary policy and equity inflows can be confirmed at a much higher level of statistical certainty. The equity flow results are also robust to the alternative specification of system GMM, reported in Column 4.

In Table 6, we provide the corresponding fund-level results for money market flows. The regression estimates show a sensitivity of money market flows to the real short rate of about 12 (based on the estimates in DGMM reported in Columns 2 and 3), compared to the corresponding estimate of about 14 for the SGMM reported in Column 4. The coefficient estimates for ΔSR are all statistically significant at the 5% level or better.

We conclude that the fund-level regressions confirm the findings of the aggregate results at the country level. The increase in statistical power due to the larger cross section and the better control for fund performance allows us to establish with greater statistical confidence that monetary policy conditions are related to economically significant investor risk shifting from fixed income to equity investment.

4.2 Causality Issues

The evidence of a strong correlation between local real interest rates and equity fund flows presented in the previous subsection can have two possible 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, 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.

Next, we examine these alternative explanations in more detail. An inflation increase—and its implied decrease of the real short rates—results from either positive aggregate demand shocks and/or negative aggregate supply shocks. Positive aggregate demand shocks increase firm profitability, which could attract net local equity fund inflows. By contrast, negative supply shocks typically generate lower output and corporate profitability. Here, positive equity fund inflows would occur in parallel to higher inflation only if local investors are contrarian equity investors. Finally, increased fiscal spending could also 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 5 and 6 should attenuate the point estimate for the real short rate and produce statistically significant point estimates for these macroeconomic measures. 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.

In Table 5, Column 5, we augment the baseline regression (DGMM2) 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 correlation coefficient between changes in the real short rate and the net equity fund flows. In particular, the three variables, ΔROA, $gGDP$, and $gGovSpd$, are all statistically insignificant, and the point estimate of ΔSR, -10.681 (t-stat $= -5.29$), is quantitatively similar to the estimate of -10.606 (t-stat $= -5.25$) for the baseline DGMM2 regression.

In Table 6, Column 5, we report the augmented regression result for money market funds. The point estimate of ΔSR is slightly reduced with the inclusion of the three additional variables, ΔROA, $gGDP$, and $gGovSpd$, but only the coefficient for $gGovSpd$ is statistically significant. Increases in government spending appear to trigger more flows into money market funds, which could indicate a Ricardian saving motive in expectation for possible higher future taxes. Yet, this effect is economically small compared to the flow effect captured by the real interest rate.

As an alternative strategy to address the aforementioned causality issues, we examine the equity flows into local funds with different investment destinations. Any covariance between corporate cash flow expectations and the local business cycle should be greatly attenuated for foreign stocks—thus allowing to distinguish between the cash flow channel as a pull factor and the risk-taking cannel as a push factor of investor flows. Columns 6 and 7 of Table 5 split the sample into $15,467$ and $58,300$ equity flow observations according to fund investment destination. The former (latter) subsample includes funds that invest more than 50% of fund assets in domestic (foreign) stocks. The point estimate for the real rate change ΔSR is -9.518 for domestically invested funds and slightly more negative at -12.241 for foreign invested funds; the difference is statistically insignificant. Therefore, disaggregation of fund flows by investment destination provides no support for the cash flow channel as the alternative explanation of the observed negative relation between equity fund flows and changes in real short rates.

The similarity in the real short rate sensitivity of equity flows across funds with a local and non-local investment focus is illustrated in Figure 4, which features the regression slope for ΔSR estimated in Table 5, Columns 6 and 7, respectively. Panel A (B) plots local investor flows into equity funds with a domestic (foreign) investment focus; the x-axis presents in both cases the country-level predicted real rate changes ΔSR spanned by the instruments reported in the table. The fund flows on the y-axis are removed of the flow component predicted by lagged fund flows ($FundFlow(-1)$ and $FundFlow(-2)$), lagged market returns ($MKT(-1)$), and lagged fund returns ($FundReturn(-1)$) as estimated in Table 5, Columns 6 and 7, respectively.[13] The slightly higher flow sensitivity for funds with a foreign investment focus is consistent with the risk-taking channel if foreign invested funds are considered slightly riskier by local investors. By contrast, the cash flow channel would predict more pronounced real rate sensitivity for domestically invested funds if the local business cycle generates a positive relation between local firm profitability and local inflation.

Figure 5 aggregates all individual fund flows in a given quarter to country averages as a function of the same predicted real short rate changes. Country averages of equity fund flows for domestically invested funds (Panel A) and foreign invested funds (Panel B) show a clear negative correlation between the two axes, with the correlation similar across the two types of funds, which feature different investment destinations and therefore do not share the same business cycle for the firms in which they invest.

We can further investigate the subset of foreign invested funds whose foreign assets are strictly confined in the eurozone.[14] The estimates reported in Table 5, Column 8 show that the coefficient of the real short rate changes for this subset of funds is even slightly higher than that for all foreign invested funds. This finding also 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 serve as an inflation hedging vehicle, this hedging benefit is absent for foreign stocks in the eurozone. Intra-union investments are undertaken at a nominally fixed exchange rate, which is by construction unrelated to the inflation differentials across member countries. A hedging motive should therefore imply a much weaker linkage between the real short rate and the equity flows into funds with a eurozone focus—a notion rejected in the data based on the result reported in Column 8.

We conclude that the equity flow effect we document in Tables 3 and 5 is unlikely to be explained by time-varying expectations about firm profitability over the local business cycle or by inflation hedging motives. Instead, the strong correlation between equity fund inflows

[13]The unadjusted flows have a very similar dispersion to the adjusted flows because the predicted flow component is relatively small.

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

and lower local real rates is likely to reflect investor risk shifting from fixed income to equity investment under loose monetary policy conditions as captured by the real short rate. In a single country setting, the real short rate is a policy choice of the central bank (via its nominal rate setting); therefore, our finding in this section suggests a powerful risk-taking channel of monetary policy.

4.3 Stock Price Effects of Monetary Policy

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 documented in the previous section.[15] Unlike the riskless rate effect, which should affect 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 monetary-policy-related asset reallocation effect. Second, we need to isolate equity fund flows induced by monetary policy conditions from all other (nonmonetary-policy-related) fund flows.

Fund returns by definition proxy for returns of those stocks that funds already heavily invest in and are likely to channel further investment into. In particular, any flow-related price pressure should be captured by fund returns. 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. We construct a *Low Investor Flow Index* ($LIFI$) based on the 15% of stocks with the lowest fund flows in each country over the previous three-year period.

Because fund flows should primarily impact the returns of the flow-sensitive stocks that funds invest in, we can identify equity flow-related price effect as the fund return in excess of the benchmark return:

$$FundReturn_{j,t} - LIFI_{c,t} = \gamma FundFlow_{j,t} + \vartheta_{j.t}. \qquad (3)$$

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

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

The second identifying step involves isolating the (predictable) fund flows induced by the cross-sectional variation in eurozone monetary policy conditions from all other fund flows represented by the residual $\kappa_{j,t}$. In the flow decomposition

$$FundFlow_{j,t} = \widehat{FundFlow}_{j,t} + \kappa_{j,t}, \tag{4}$$

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

$$\widehat{FundFlow}_{j,t} = \alpha_1 \Delta SR_{c,t} + \alpha_2 \widehat{FundFlow}_{j,t-1} + \alpha_3 \widehat{FundFlow}_{j,t-2} + \mu_j, \tag{5}$$

where the coefficients α_1, α_2, and α_3 correspond to the estimates obtained in Eq.(2). To derive the predicted fund flows strictly from changes in short-term real interest rates, we drop the market returns and fund returns from the equation. Similarly, we can further relate $\widehat{FundFlow}_{j,t-1}$ and $\widehat{FundFlow}_{j,t-2}$ to lagged changes of short-term real interest rates. Substitution into Eq.(4) and Eq.(3) yields the specification

$$FundReturn_{j,t} - LIFI_{c,t} = \beta_0 + \beta_1 \Delta SR_{c,t} + \beta_2 \Delta SR_{c,t-1} + \beta_3 \Delta SR_{c,t-2} + \nu_j + \varepsilon_{j.t}, \tag{6}$$

with linear constraints $\beta_1 = \gamma\alpha_1$, $\beta_2 = \gamma\alpha_1\alpha_2$, and $\beta_3 = \gamma\alpha_1(\alpha_2^2 + \alpha_3)$, and small terms in $\Delta SR_{c,t-k}$ with $k > 2$ ignored. Eq.(6) can be estimated simultaneously with Eq.(5) under the two constraints, $\beta_2 = \alpha_2\beta_1$ and $\beta_3 = (\alpha_2^2 + \alpha_3)\beta_1$. The sum of the constrained coefficients, β_1, β_2, and β_3, 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.

Table 7 provides the estimation results for the two equations (5) and (6) with fund returns benchmarked against the $LIFI$ index. In Columns 1–3, we report regressions in which each country has the same regression weight to best use the full variation in the real short rates. Because the number of funds, $N(c)$, varies substantially from 76 in Portugal to $2,385$ in France, an equal fund weight would effectively limit our empirical inference to the policy variations of the three largest countries, France, Germany, and Italy, which combined represent about 75% of all fund observations. By contrast, an equal country weight implies that each fund observation is weighted by $[1/8] \times [1/N(c)]$. Another consideration with respect to regression weights concerns the relative importance of local investors in various countries. The share of the local capital market held by local institutional investors, $LocInstShare(c)$, varies from 1.1% in Austria to 10.7% in Germany. Accordingly, we expect the fund flows identified in Eq.(5) to have a significantly larger price impact in Germany than in Austria. In Columns 4–6, we scale the country weights by $LocInstShare(c)$. This puts more weight on fund flows in locations where local institutional investors matter most and should increase the estimated coefficients in the excess return equation, (6).

In Table 7, specifications 1 and 4 feature no fixed effects for the second equation, whereas country fixed effects are used in specifications 2 and 5 and fund fixed effects in specifications 3 and 6. Estimation of the first equation is undertaken in first differences similar to the DGMM estimates reported in Table 5, Columns 2–3. When equal country weights are used, the simultaneous equation yields autocorrelation estimates of 0.24 and 0.06 for $Fundflow(-1)$ and $Fundflow(-2)$, respectively. The corresponding coefficient for changes in real short rates, ΔSR, is -10.4, slightly smaller than the previous single-equation estimate of -9.8 (in DGMM1 of Table 5). Overall, the coefficient estimates in the first equation are similar across all specifications, 1–6.

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 of ΔSR consists in the sum $\widehat{\beta}_1 + \widehat{\beta}_2 + \widehat{\beta}_3$. Under equal country weights in Columns 1–3, the total return effect of ΔSR is approximately $\widehat{\beta}_1 + \widehat{\beta}_2 + \widehat{\beta}_3 \approx -14$, implying that a 10-basis-point decrease in the short-term real interest rate increases the relative valuation of flow-sensitive stocks by roughly 1.4%. By contrast, $LocInstShare(c)$-adjusted country weights reported in Columns 4–6 imply a total excess return effect more than twice as large, with $\widehat{\beta}_1 + \widehat{\beta}_2 + \widehat{\beta}_3 \approx -34$. This means 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. Conversely, if the home bias is small, an accommodating monetary policy is likely to spread asset price inflation worldwide.

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.

4.4 Robustness

We undertake a variety of robustness checks. First, in light of the concern that some exceptional monetary measures undertaken during the financial crisis may taint our inference based on the real short rate, we repeat our analysis for the pre-crisis period covering 2003–2007/q2. Table 8 reports the regression results for equity funds in Panel A and money market funds in Panel B. We find that our evidence is qualitatively robust to this modified period. The estimate of ΔSR for equity funds is about -11 for this period, which is comparable to the estimate of about -10 for the full period. A decrease in the real short rate is also accompanied by substantial money market outflows during this period. Thus, our results are not driven by the crisis period.

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 learn about the realized inflation with a quarterly publication delay. Yet, we can go one step further and estimate the expected real rate change more precisely based on data from national consumer surveys. Our result suggests that national consumer surveys on inflation expectations have some additional (small) explanatory power over the quarterly realized inflation beyond lagged inflation.[16] In Table 9, Panels A and B, we replace the real rate change with the expected real rate changes [ΔSR (expected)] to explain flows for equity and money market funds, respectively. Compared to Tables 3, Panel A and Table 4, Panel A, the statistical significance of the expected real rate effect is slightly stronger but remains qualitatively similar.

The third robustness test concerns the weights used for the disaggregate flow regressions. We replace the fund-value weights used in Tables 5 and 6 with equal fund weights and discard the very small funds with a total net asset value of less than U.S. \$10 million. Such equal-weighted flow regressions produce very similar point estimates for the effect of changes in the

[16] 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.

real short rate on equity and money market flows. Columns 1–2 of Table 10 report the DGMM2 estimates. The point estimate for ΔSR is -9.507 for equity funds and 11.268 for money market funds under the equal-weighted approach, compared to -10.606 and 11.540 under the value-weighted approach. We conclude that the interest rate effect on fund flows does not depend on fund size.

The fourth robustness test uses alternative monetary policy rates in the disaggregate fund flow regressions of Tables 5 and 6. Columns 3–4 of Table 10 replace ΔSR with $\Delta Euribor$ (changes in the three-month real euro interbank offered rate), and Columns 5–6 replace ΔSR with ΔTR (changes in the Taylor rule residual). The results are qualitatively very similar to those reported in Tables 5 and 6. The DGMM2 estimates for $\Delta Euribor$ (ΔTR) are 10.6 and -11.533 (-15.481 and 15.594), respectively, for equity funds and money market funds, compared to the corresponding estimates of -10.606 and 11.540 for ΔSR. The numerically larger point estimates for the ΔTR coefficient reflect the fact that the standard deviation of the Taylor rule residual changes is about 24% smaller than the standard deviation of the real short rate changes. The disaggregate fund flow results are therefore robust to the two alternative measures of the monetary policy rate.

The fifth robustness test uses an alternative benchmark return index. Rather than constructing the benchmark index based on fund flows, we construct for each country a value-weighted *Low Fund Holding Index* ($LFHI$), which comprises 15% of local stocks with the lowest share of fund investment overall. The $LFHI$ index generally behaves similarly to the $LIFI$ index, with an overall return correlation of 0.98 between the two indices. We repeat the simultaneous equation regressions of Table 7 using this alternative index as the relevant return benchmark. We find similar results. To conserve space, Columns 7–8 of Table 10 report only the second equation of the simultaneous equation system. Equal country weights imply a total stock price effect of $\widehat{\beta}_1 + \widehat{\beta}_2 + \widehat{\beta}_3 \approx -12$, whereas $LocInstShare(c)$-adjusted country weights imply a total effect of $\widehat{\beta}_1 + \widehat{\beta}_2 + \widehat{\beta}_3 \approx -36$. Overall, using either low fund holdings or low fund flows to proxy for the 'non-investability' of a stock gives quantitatively similar estimates of the stock price effect.

The last robustness test concerns the alternative threshold for constructing the $LIFI$. Table 7 constructs the value-weighted $LIFI$ index using the 15% stocks in each country with the least inflow and outflow of fund investors during the past three years. Columns 9–10 use an alternative threshold of 10%, and Columns 11–12 use 20%. Overall, the quantitative return results of Table 7 become slightly stronger for the 10% threshold and slightly weaker for the more inclusive 20% cut-off, but the results remain qualitatively robust to the alternative thresholds.

5 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. Local equity fund inflows can reveal a change in investment preferences by local investors triggered either by the increased opportunity cost of riskless investment (under low real rates) or an increased investor risk tolerance.

First, we find that loose local monetary policy conditions (measured by decrease in either the real short-term interest rate or the Taylor rule residual) 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. This evidence is obtained in both the aggregate country-level analysis as well as the (more powerful) fund-level analysis. The difference between the highest and lowest real short rate (across the eight eurozone sample countries) was on average 53 basis points. Based on our regression estimates, a half-percentage-point lower real short rate is associated with a 4% incremental money market outflow and a 5% incremental equity market inflow relative to fund assets under management.

Second, we examine whether investor asset allocation decisions respond directly to the local business cycle by using local aggregate income growth, government spending growth, and changes in firm profitability as control variables for the regression of investor fund flows on changes in the local real short rate. None of these business cycle proxies features any statistical significance for investor asset allocation decisions, whereas changes in the local real rate retain their full economic and statistical significance. Furthermore, disaggregate fund flows allow us to distinguish flows into those funds with a primarily non-local (i.e., foreign or eurozone equity) investment focus. For those funds, the local business cycle is unlikely to generate any systematic variations in the expected cash flows of the invested firms. Yet, those fund flows show the same sensitivity to the local real rate as the overall local flows. Changing firm cash flow expectations related to the local business cycle cannot easily account for this identical real rate sensitivity of fund flows irrespective of a fund's investment destination; whereas investor sensitivity to the local real rate predicts a similar flows dynamics across funds with different investment destinations.

Third, we explore whether the asset reallocation process explained by local monetary policy conditions contributes to equity price inflation. To this end, we identify in each country the return difference between the stocks held by local equity funds and a control group of stocks least prone to fund flows. A structural simultaneous equation approach allows us to assert that the investor asset reallocation toward equity funds triggered by loose local monetary policy conditions generates stock price inflation relative to a benchmark group of stocks with low 'investability.' The observed excess return in investable stocks is largest 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 subject to local sentiments about the real short rate 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 a powerful link between monetary policy and investors' asset allocation decisions. Loose monetary policy appears to increase investor risk tolerance and thereby contribute to investor risk taking through increased equity investment; asset price inflation is indicative of such endogenous risk tolerance. In practice, it is often difficult to identify the monetary policy component of asset price inflation, partly due to the high overall stock market volatility. We argue that knowledge about investors' asset allocation decisions can serve as a useful complementary source of information on 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 bears implications on 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 (2013) argues that currency pegs (such as the gold standard or more recently the common currency in the eurozone) augment variations in the 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 the 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.	Datasteam
$gGDP$	Quarterly growth of real GDP.	Datastream
INF	Quarterly inflation rate.	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
SR	Quarterly short-term real interest rate, calculated as the difference between EOINA and the quarterly inflation rate.	Datastream
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
$Euribor$	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
MKT	Quarterly return on the MSCI country market index.	Datastream

Appendix A continued.

Variable	Description	Source
$LIFI$	Quarterly return on the value-weighted index of the 15% local stocks with the lowest absolute fund inflows and outflows over the previous three years; fund flows are measured by the change in the aggregate share holdings of all funds relative to a stock's shares outstanding.	Thomson Financial and Datastream
$LFHI$	Quarterly return on the value-weighted index of the 15% of stocks with the lowest average fund holdings overall. Fund holdings are aggregated across all funds and scaled by a stock's shares outstanding.	Thomson Financial and Datastream
$FundReturn$	Net quarterly return of a fund.	Lipper
TNA	Total net asset value of a fund.	Lipper
Disaggregate $FundFlow$	A fund's net quarterly flow, calculated as its net dollar flow scaled by the beginning-of-period TNA. The net dollar flow is estimated by the difference between the end-of-period TNA and the product of the beginning-of-period TNA and one plus the current fund return.	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.	Lipper
$LocInstShare$	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 (2012).	Bartram, Griffin, and Ng (2012)

Appendix A continued.

Variable	Description	Source
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

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Table 1: Summary Statistics of 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 from 2003/1–2010/4. We also report the short-term real interest rate (SR) and Taylor rule residual (TR) by country as well as their cross-country averages. The cross-country averages of changes in the short-term real interest rate (ΔSR) and changes in the Taylor rule residual (ΔTR) 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 $\times 100$						
$EONIA$	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) $\times 100$						
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
Taylor Rule Residual (TR) $\times 100$						
Austria	32	−0.002	0.035	0.248	−0.497	0.417
Finland	32	0.076	0.153	0.203	−0.551	0.324
France	32	0.026	−0.012	0.254	−0.438	0.528
Germany	32	0.054	0.056	0.220	−0.361	0.362
Italy	32	−0.030	−0.060	0.239	−0.406	0.475
Netherlands	32	0.041	0.025	0.266	−0.400	0.516
Portugal	32	−0.034	−0.111	0.234	−0.492	0.381
Spain	32	−0.132	−0.188	0.222	−0.525	0.347
All TR	256	0.000	−0.002	0.241	−0.551	0.528
All ΔTR	248	−0.015	−0.006	0.089	−0.362	0.257

Table 2: Summary Statistics of Equity and Money Market Funds

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 from 2003/1–2010/4. Also reported are the net equity and money market flows at the fund level, fund returns ($FundReturn$), and fund size (TNA) in million U.S. dollars. We calculate a fund's net quarterly flow as its net dollar flow scaled by the beginning-of-period TNA. The net dollar flow is estimated by $[TNA_t - TNA_{t-1} \times (1 + FundReturn_t)]$. The aggregate fund flow is the aggregate net dollar flow for all funds in a country scaled by their aggregate beginning-of-period TNA. The last two rows of the table report the MSCI country market index return (MKT) and the value-weighted index return for the 15% of stocks with the lowest absolute fund inflows and outflows measured over the previous three year period ($LIFI$).

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 Fund Characteristics						
Disaggregate Fund Flows	89,415	0.008	−0.011	0.161	−0.751	6.619
Fund Return	89,750	0.023	0.019	0.115	−0.565	0.602
Fund Size (TNA)	89,750	104.512	30.405	249.043	< 0.001	7791.410
Money Market Fund Characteristics						
Disaggregate Fund Flows	24,932	−0.015	−0.027	0.166	−0.820	6.539
Fund Return	24,950	0.011	0.010	0.054	−0.578	0.275
Fund Size (TNA)	24,950	574.025	125.505	1522.492	< 0.001	25000.000
Equity Index Returns						
MKT	256	0.034	0.031	0.142	−0.432	0.388
$LIFI$	256	0.036	0.029	0.135	−0.382	0.442

Table 3: Aggregate Equity Fund Flows and Innovations to Monetary Policy Rates

Reported are the regression results for the quarterly country aggregate net inflows into equity funds domiciled and marketed exclusively in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period from 2003/1–2010/4. Panels A and B use the short-term real interest rates and the Taylor rule residuals, respectively, as measures for local monetary policy conditions. 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 4–5 provide the estimates using the difference generalized method of moments (DGMM) and system generalized method of moments (SGMM), respectively. Columns 6–10 report the corresponding results with the monetary policy rate proxied by the Taylor rule residual. Changes (from the previous quarter) in the short-term real interest rates and the Taylor rule residuals are denoted by ΔSR and ΔTR, respectively; $FundFlow(-1)$ denotes the fund flow in the previous quarter; $MKT(-1)$ is the country stock market return in the previous quarter. 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.

Panel A: Short-Term Real Interest Rates

Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)
ΔSR	−4.538	−9.638	−9.651	−9.261	−9.663
	[−2.04]	[−3.90]	[−3.96]	[−4.91]	[−5.21]
$FundFlow(-1)$	0.332	0.298	0.315	0.285	0.348
	[4.19]	[2.55]	[2.75]	[2.52]	[3.37]
$MKT(-1)$	−0.052	−0.043	−0.052	−0.039	−0.053
	[−1.62]	[−1.40]	[−1.52]	[−1.23]	[−1.66]
Obs.	254	246	246	254	254
Adj.R^2	0.297				
Instruments					
ΔSR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
$Total\ Number$		6	9	9	12
$AR(1)$		0.012	0.010	0.012	0.010
$AR(2)$		0.545	0.546	0.560	0.506
Hansen Test		0.197	0.372	0.537	0.728

Panel B: Taylor Rule Residuals

Dep. Variable: Fund Flow	LSDV (6)	DGMM1 (7)	DGMM2 (8)	SGMM1 (9)	SGMM2 (10)
ΔTR	−6.229	−14.085	−14.074	−15.035	−15.781
	[−1.81]	[−3.90]	[−3.96]	[−4.78]	[−5.25]
$FundFlow(-1)$	0.358	0.298	0.318	0.252	0.323
	[4.25]	[2.39]	[2.65]	[1.89]	[2.76]
$MKT(-1)$	−0.043	−0.028	−0.038	−0.022	−0.037
	[−1.37]	[−0.76]	[−0.98]	[−0.57]	[−1.01]
Obs.	248	240	240	248	248
Adj.R^2	0.312				
Instruments					
ΔTR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
$Total\ Number$		6	9	9	12
$AR(1)$		0.021	0.018	0.026	0.018
$AR(2)$		0.342	0.349	0.396	0.362
Hansen Test		0.166	0.348	0.528	0.697

Table 4: Aggregate Money Market Fund Flows and Innovations to Monetary Policy Rates

Reported are the regression results for the quarterly country aggregate net inflows into money market funds domiciled and marketed exclusively in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period from 2003/1–2010/4. Panels A and B use the short-term real interest rates and the Taylor rule residuals, respectively, as measures for local monetary policy conditions. 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 provide the estimates using the difference generalized method of moments (DGMM) and system generalized method of moments (SGMM), respectively. Columns 6–10 report the corresponding results with the monetary policy rate proxied by the Taylor rule residual. Changes (from the previous quarter) in the short-term real interest rate and changes in the Taylor rule residuals are denoted by ΔSR and ΔTR, respectively; $FundFlow(-1)$ denotes the fund flow in the previous quarter; $MKT(-1)$ is the country stock market return in the previous quarter. 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.

Panel A: Short-Term Real Interest Rates					
Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)
ΔSR	7.711	8.503	7.766	9.090	8.948
	[2.18]	[1.91]	[1.63]	[2.54]	[2.60]
$FundFlow(-1)$	0.363	0.361	0.315	0.372	0.326
	[5.06]	[5.19]	[4.91]	[6.53]	[5.89]
$MKT(-1)$	0.065	−0.009	−0.006	0.009	0.011
	[0.91]	[−0.10]	[−0.07]	[0.11]	[0.12]
Obs.	249	240	240	249	249
$Adj.R^2$	0.228				
Instruments					
ΔSR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
Total Number		6	9	9	12
AR(1)		0.011	0.009	0.008	0.007
AR(2)		0.801	0.870	0.897	0.974
Hansen Test		0.360	0.330	0.730	0.579

Panel B: Taylor Rule Residuals					
Dep. Variable: Fund Flow	LSDV (6)	DGMM1 (7)	DGMM2 (8)	SGMM1 (9)	SGMM2 (10)
ΔTR	11.947	12.523	11.248	11.895	11.487
	[2.30]	[1.84]	[1.55]	[1.82]	[1.73]
$FundFlow(-1)$	0.361	0.363	0.315	0.365	0.321
	[5.00]	[5.20]	[4.90]	[5.97]	[5.04]
$MKT(-1)$	0.040	−0.011	−0.008	−0.005	−0.003
	[0.57]	[−0.12]	[−0.08]	[−0.05]	[−0.04]
Obs.	244	235	235	244	244
$Adj.R^2$	0.237				
Instruments					
ΔTR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
Total Number		6	9	9	12
$AR(1)$		0.011	0.009	0.011	0.008
$AR(2)$		0.802	0.872	0.800	0.870
Hansen Test		0.318	0.342	0.714	0.709

Table 5: Disaggregate Equity Fund Flows and Innovations to Monetary Policy Rates

Reported are the regression results for the quarterly net inflows into each equity fund domiciled and marketed in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period from 2003/1–2010/4. Each country-quarter is given the same weight and each fund within a country is weighted by fund size at the beginning of the period. To eliminate the need for time fixed effects, all variables are expressed as deviations from their cross-sectional means. The regressors are (i) changes in the short-term real interest rate ΔSR; (ii) fund flows at lags 1 and 2 given by $FundFlow(-1)$ and $FundFlow(-2)$, respectively; (iii) the country stock market return in the previous quarter $MKT(-1)$; (iv) individual fund returns in the previous quarter given by $FundReturn(-1)$; (v) changes in aggregate corporate profitability, proxied by changes in return on assets (ΔROA) at the country-level; (vi) GDP growth ($gGDP$); and (vii) growth in real government expenditure ($gGovSpd$). Column 1 states the result for the least square dummy variable (LSDV) regression without instruments. Columns 2 and 3 provide the estimates using the difference generalized method of moments (DGMM) estimator, whereas Column 4 reports estimates based on the system generalized method of moments (SGMM). Column 5 uses the same setup as Column 3 but includes three additional regressors, ΔROA, $gGDP$, and $gGovSpd$. Column 6 provides the DGMM estimate for the subsample of funds that invest more than 50% of their fund assets in domestic stocks. Column 7 provides the DGMM estimate for the subsample of funds that invest more than 50% of their fund assets in foreign stocks; column 8 further restricts the foreign stocks to be strictly inside the eurozone. 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.

	Full Fund Sample					Funds with Specific Investment Focus		
						Domestic	Foreign	Eurozone
Dep. Variable:	LSDV	DGMM1	DGMM2	SGMM	DGMM3	DGMM4	DGMM5	DGMM6
Fund Flow	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔSR	−3.694	−9.789	−10.606	−7.325	−10.681	−9.518	−12.212	−16.147
	[−3.52]	[−5.00]	[−5.25]	[−5.03]	[−5.29]	[−2.15]	[−4.93]	[−3.24]
$FundFlow(-1)$	0.195	0.349	0.340	0.345	0.340	0.275	0.352	0.207
	[12.27]	[14.54]	[14.30]	[14.25]	[14.31]	[6.82]	[13.31]	[5.76]
$FundFlow(-2)$	0.061	0.128	0.129	0.159	0.130	0.187	0.117	0.038
	[5.70]	[4.77]	[4.85]	[5.74]	[4.86]	[2.68]	[4.26]	[0.65]
$MKT(-1)$	−0.023	0.017	−0.023	−0.035	−0.020	−0.632	−0.011	0.072
	[−1.10]	[0.54]	[−0.76]	[−1.19]	[−0.66]	[−1.86]	[−0.33]	[1.12]
$FundReturn(-1)$			0.197	0.158	0.198	0.401	0.249	0.342
			[5.36]	[5.97]	[5.39]	[1.89]	[5.60]	[1.57]
ΔROA					−0.137			
					[−0.62]			
$gGDP$					−0.072			
					[−0.25]			
$gGovSpd$					0.015			
					[1.24]			
Obs.	78,735	73,767	73,767	78,735	73,767	15,467	58,300	10,398
$Adj.R^2$	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 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3	Lags 3-4	Lags 2-3	Lags 2-3
$FundReturn$			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	8	12	11	8	8	8
AR(1)		0.000	0.000	0.000	0.000	0.000	0.000	0.000
AR(2)		0.714	0.708	0.181	0.745	0.165	0.833	0.565
Hansen Test		0.057	0.114	0.002	0.087	0.215	0.465	0.314

Table 6: Disaggregate Money Market Fund Flows and Innovations to Monetary Policy Rates

Reported are the regression results for the quarterly net inflows into each money market fund domiciled and marketed in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period from 2003/1–2010/4. Similar to the set-up in Table 5, each country-quarter is given the same weight and each fund within a country is weighted by fund size at the beginning of the period. The regressors and the instrument set used are the same as those in Columns 1–5 of Table 5.

Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM (4)	DGMM3 (5)
ΔSR	12.218	11.745	11.540	13.537	11.468
	[3.34]	[2.29]	[2.27]	[3.78]	[2.21]
$FundFlow(-1)$	0.149	0.290	0.284	0.286	0.282
	[3.26]	[3.17]	[3.13]	[3.26]	[3.08]
$FundFlow(-2)$	0.006	0.103	0.101	0.107	0.100
	[0.18]	[1.85]	[1.82]	[2.05]	[1.81]
$MKT(-1)$	0.004	−0.003	−0.008	−0.003	−0.007
	[0.07]	[−0.04]	[−0.12]	[−0.05]	[−0.10]
$FundReturn(-1)$			0.952	0.684	0.932
			[1.66]	[1.74]	[1.63]
ΔROA					0.250
					[0.67]
$gGDP$					−0.079
					[−0.10]
$gGovSpd$					0.072
					[2.58]
Obs.	19,694	17,659	17,659	19,694	17,659
Adj.R^2	0.112				
Instruments					
ΔSR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
$FundReturn$			Lags 2-3	Lags 2-3	Lags 2-3
ΔROA					Lag 0
$gGDP$					Lag 0
$gGovSpd$					Lag 0
Total Number		6	8	12	11
AR(1)		0.000	0.000	0.000	0.000
AR(2)		0.239	0.240	0.312	0.227
Hansen Test		0.724	0.899	0.801	0.938

Table 7: 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 short-term real interest rates (ΔSR) and is estimated (as before) using the DGMM approach. The second equation relates fund excess returns, $FundReturn_{j,t} - LIFI_{c,t}$, given in Eq. (6) to contemporaneous and lagged short-term real interest rate changes with cross-equation restrictions 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, country fixed effects, or fund fixed effects. To eliminate the need for time fixed effects, all variables are expressed as deviations from their cross-sectional means. The sample covers all locally distributed and marketed equity funds (with a total net asset value of U.S. $10 million or more at the beginning of the period) in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 2003/1–2010/4. Columns 1–3 present results based on equal country weights. Each of the $N(c)$ local funds in country c carries the same regression weight $[1/8] \times [1/N(c)]$ each quarter. Columns 4–6 use country weights given by $LocInstShare(c)$, defined as the proportion of the local stock market held by local institutional investors. Thus, each fund has a regression weight of $[LocInstShare(c)/\sum_c LocInstShare(c)] \times [1/N(c)]$ 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.

	Equal Country Weights			$LocInstShare$ as Country Weight		
	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Variable Equation 1: $FundFlow_{j,t}$						
ΔSR	−10.427	−10.572	−10.427	−10.511	−10.763	−10.494
	[−9.71]	[−9.92]	[−9.72]	[−9.62]	[−9.94]	[−9.62]
$FundFlow(-1)$	0.238	0.237	0.237	0.228	0.229	0.226
	[25.94]	[26.04]	[25.82]	[25.14]	[25.40]	[24.97]
$FundFlow(-2)$	0.063	0.063	0.063	0.048	0.048	0.047
	[12.16]	[12.14]	[12.05]	[9.34]	[9.44]	[9.16]
Dep. Variable Equation 2: $FundReturn_{j,t} - LIFI_{c,t}$						
ΔSR	−10.051	−10.142	−10.512	−24.952	−25.124	−25.415
	[−15.20]	[−15.27]	[−16.06]	[−32.93]	[−33.15]	[−33.82]
$\Delta SR(-1)$	−2.394	−2.404	−2.500	−5.737	−5.764	−5.839
	[−15.20]	[−15.27]	[−16.06]	[−32.93]	[−33.15]	[−33.82]
$\Delta SR(-2)$	−1.206	−1.207	−1.257	−2.534	−2.541	−2.577
	[−15.20]	[−15.27]	[−16.06]	[−32.93]	[−33.15]	[−33.82]
Sum of ΔSR Coefficients	−13.651	−13.753	−14.269	−33.223	−33.429	−33.831
Country Fixed Effects	NO	YES	NO	NO	YES	NO
Fund Fixed Effects	NO	NO	YES	NO	NO	YES
$Obs.$	57,697	57,697	57,697	57,697	57,697	57,697
Instruments (Eq.1 and Eq. 2)						
ΔSR	Lags 1-3	Lags 1-3	Lags 1-3	Lags 1-3	Lags 1-3	Lags 1-3
$FundFlow$	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

Table 8: Sub-Sample Analysis: 2003–2007/q2

Reported are the regression results for the quarterly country aggregate net inflows into equity funds domiciled and marketed exclusively in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period from the beginning of 2003 to the end of the second quarter of 2007. Panels A and B report the results for equity funds and money market funds, respectively. 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 4–5 provide the estimates using difference generalized method of moments (DGMM) and system generalized method of moments (SGMM), respectively. Columns 6–10 report the corresponding results for money market funds. Changes (from the previous quarter) in the short-term real interest rate are denoted by ΔSR; $FundFlow(-1)$ denotes the fund flow in the previous quarter; $MKT(-1)$ is the country stock market return in the previous quarter. 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.

Panel A: Equity Funds

Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)
ΔSR	−6.946	−10.962	−11.054	−11.143	−11.459
	[−2.27]	[−2.85]	[−2.91]	[−3.00]	[−3.18]
$FundFlow(-1)$	0.151	0.067	0.030	0.162	0.129
	[1.68]	[0.76]	[0.32]	[2.54]	[2.03]
$MKT(-1)$	−0.081	−0.054	−0.048	−0.061	−0.065
	[−1.75]	[−1.32]	[−1.32]	[−1.14]	[−1.25]
Obs.	142	134	134	142	142
$Adj.R^2$	0.350				
Instruments					
ΔSR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
$Total\ Number$		6	9	9	12
$AR(1)$		0.025	0.032	0.010	0.010
$AR(2)$		0.112	0.102	0.170	0.155
Hansen Test		0.612	0.748	0.540	0.823

Panel B: Money Market Funds

Dep. Variable: Fund Flow	LSDV (6)	DGMM1 (7)	DGMM2 (8)	SGMM1 (9)	SGMM2 (10)
ΔSR	6.177	9.979	13.274	12.378	14.380
	[1.08]	[1.05]	[1.43]	[2.02]	[2.56]
$FundFlow(-1)$	0.162	0.289	0.212	0.345	0.268
	[1.25]	[2.26]	[1.82]	[3.32]	[3.15]
$MKT(-1)$	0.004	0.017	0.012	−0.004	−0.012
	[0.04]	[0.13]	[0.10]	[−0.04]	[−0.10]
Obs.	137	128	128	137	137
$Adj.R^2$	0.283				
Instruments					
ΔSR		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
$Total\ Number$		6	9	9	12
$AR(1)$		0.020	0.020	0.011	0.009
$AR(2)$		0.681	0.596	0.610	0.538
Hansen Test		0.220	0.417	0.466	0.908

Table 9: Expected Real Short Rate and Fund Flows

Reported are the regression results for the quarterly country aggregate net inflows into equity funds domiciled and marketed exclusively in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period from the beginning of 2003 to the end of 2010. Panels A and B report the results for equity funds and money market funds, respectively. 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 4–5 provide the estimates using difference generalized method of moments (DGMM) and system generalized method of moments (SGMM), respectively. Columns 6–10 report the corresponding results for money market funds. SR (*expected*) is the difference between the quarterly EONIA and the quarterly expected inflation rate derived from the European Commission's Consumer Survey data. ΔSR (*expected*) denotes the quarterly change of SR (*expected*). $FundFlow(-1)$ denotes the fund flow in the previous quarter; $MKT(-1)$ is the country stock market return in the previous quarter. 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.

Panel A: Equity Funds

Dep. Variable: Fund Flow	LSDV (1)	DGMM1 (2)	DGMM2 (3)	SGMM1 (4)	SGMM2 (5)
ΔSR (*expected*)	−5.398	−12.390	−12.275	−10.954	−11.292
	[−2.01]	[−5.09]	[−5.07]	[−5.76]	[−6.60]
$FundFlow(-1)$	0.334	0.310	0.324	0.298	0.351
	[4.21]	[2.68]	[2.82]	[2.64]	[3.38]
$MKT(-1)$	−0.051	−0.039	−0.047	−0.037	−0.050
	[−1.58]	[−1.18]	[−1.31]	[−1.15]	[−1.52]
Obs.	254	246	246	254	254
$Adj.R^2$	0.296				
Instruments					
ΔSR (*expected*)		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
Total Number		6	9	9	12
$AR(1)$		0.011	0.010	0.011	0.009
$AR(2)$		0.442	0.445	0.475	0.438
Hansen Test		0.203	0.400	0.568	0.712

Panel B: Money Market Funds

Dep. Variable: Fund Flow	LSDV (6)	DGMM1 (7)	DGMM2 (8)	SGMM1 (9)	SGMM2 (10)
ΔSR (*expected*)	9.683	13.428	12.596	12.713	12.563
	[2.26]	[2.38]	[2.19]	[2.77]	[2.78]
$FundFlow(-1)$	0.364	0.365	0.319	0.373	0.324
	[5.07]	[5.06]	[4.98]	[6.66]	[5.86]
$MKT(-1)$	0.062	−0.018	−0.014	0.004	0.005
	[0.87]	[−0.19]	[−0.15]	[0.04]	[0.06]
Obs.	249	240	240	249	249
$Adj.R^2$	0.228				
Instruments					
ΔSR (*expected*)		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
MKT		Lags 2-3	Lags 2-3	Lags 2-3	Lags 2-3
Total Number		6	9	9	12
$AR(1)$		0.011	0.008	0.008	0.007
$AR(2)$		0.808	0.875	0.912	0.993
Hansen Test		0.394	0.326	0.747	0.567

Table 10: Robustness

Panel A repeats DGMM2 of Tables 5 and 6 with alternative monetary policy rates. Columns 1–2 use the same real short rate changes (ΔSR) as those in Tables 5 and 6 but employ an equal-fund-weight approach. Under this approach, each country-quarter is given the same weight, and the weight is then equally divided among all fund observations in a country; very small funds with a total net asset value of less than U.S. $10 million are excluded. Columns 3–4 replace ΔSR of Tables 5 and 6 with the 3-month real Euribor changes ($\Delta Euribor$), and Columns 5–6 replace ΔSR with the changes in the Taylor rule residual (ΔTR). The estimates for equity funds are reported in Columns 1, 3, and 5, and money market funds in Columns 2, 4, and 6. Panel B repeats the simultaneous equation estimation of Table 7 with alternative benchmark return indexes. Columns 7–8 measure fund excess returns against the *low fund holdings index* ($LFHI$), which is a value-weighted return index for the 15% of stocks in each country with the lowest average fund holdings. Columns 9–10 use an alternative $LIFI$ index constructed using the 10% least traded stocks in the past three years, and Columns 11–12 use an alternative $LIFI$ index constructed using the 20% least traded stocks. As in Table 7, we report both equal-country-weight (Columns 7, 9, and 11) and $LocInstShare$ weighted (Columns 8, 10, and 12) regression results. To conserve space, we only report Equation 2 of the simultaneous equation system.

Panel A: Alternative Monetary Policy Rate

	ΔSR−Equal Weight		$\Delta Euribor$		ΔTR	
	(1)	(2)	(3)	(4)	(5)	(6)
	Equity	M.M.	Equity	M.M.	Equity	M.M.
ΔSR, $\Delta Euribor$, or ΔTR	−9.507	11.268	−10.600	11.533	−15.481	15.594
	[−5.81]	[1.93]	[−5.25]	[2.27]	[−5.09]	[2.12]
$FundFlow(-1)$	0.242	0.255	0.340	0.284	0.340	0.285
	[11.55]	[5.47]	[14.30]	[3.13]	[14.32]	[3.13]
$FundFlow(-2)$	0.072	0.096	0.129	0.101	0.129	0.101
	[4.67]	[2.43]	[4.85]	[1.82]	[4.86]	[1.82]
$MKT(-1)$	−0.020	−0.002	−0.023	−0.008	−0.024	−0.004
	[−0.79]	[−0.02]	[−0.76]	[−0.12]	[−0.80]	[−0.06]
$FundReturn(-1)$	0.182	0.487	0.197	0.952	0.194	0.961
	[6.72]	[1.70]	[5.36]	[1.66]	[5.29]	[1.67]
Obs.	57,697	16,262	73,767	17,659	73,767	17,659

Panel B: Alternative Benchmark Return Index for Equation 2 of the Simultaneous Equation System

	$LHFI$-15% cutoff		$LIFI$-10% cutoff		$LIFI$-20% cutoff	
	(7)	(8)	(9)	(10)	(11)	(12)
	Equal Wt.	*LocInst.*	Equal Wt.	*LocInst.*	Equal Wt.	*LocInst.*
ΔSR	−9.075	−26.972	−10.827	−29.368	−7.991	−16.523
	[−13.29]	[−34.33]	[−14.40]	[−34.74]	[−13.30]	[−24.90]
$\Delta SR(-1)$	−2.228	−6.191	−2.575	−6.747	−1.900	−3.796
	[−13.29]	[−34.33]	[−14.40]	[−34.74]	[−13.30]	[−24.90]
$\Delta SR(-2)$	−1.172	−2.732	−1.295	−2.978	−0.956	−1.675
	[−13.29]	[−34.33]	[−14.40]	[−34.74]	[−13.30]	[−24.90]
Sum of ΔSR Coeffients	−12.475	−35.895	−14.697	−39.093	−10.847	−21.994
Fund Fixed Effects	YES	YES	YES	YES	YES	YES
Obs.	57,697	57,697	57,697	57,697	57,697	57,697

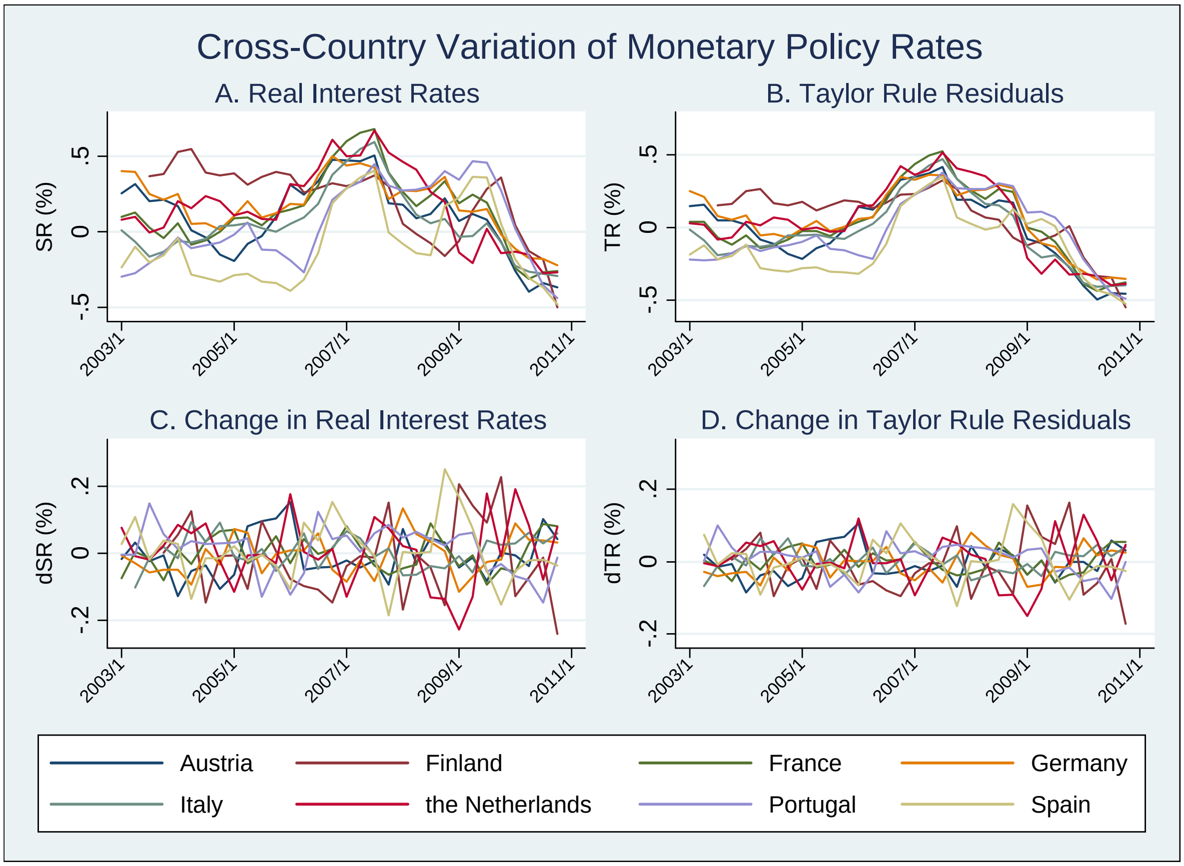

Figure 1: Plotted in Panels A and B are the quarterly short-term real interest rate (SR) and Taylor rule residuals (TR), respectively, for each of the eight eurozone countries–Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain from 2003/1–2010/4. Panels C and D plot the quarterly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

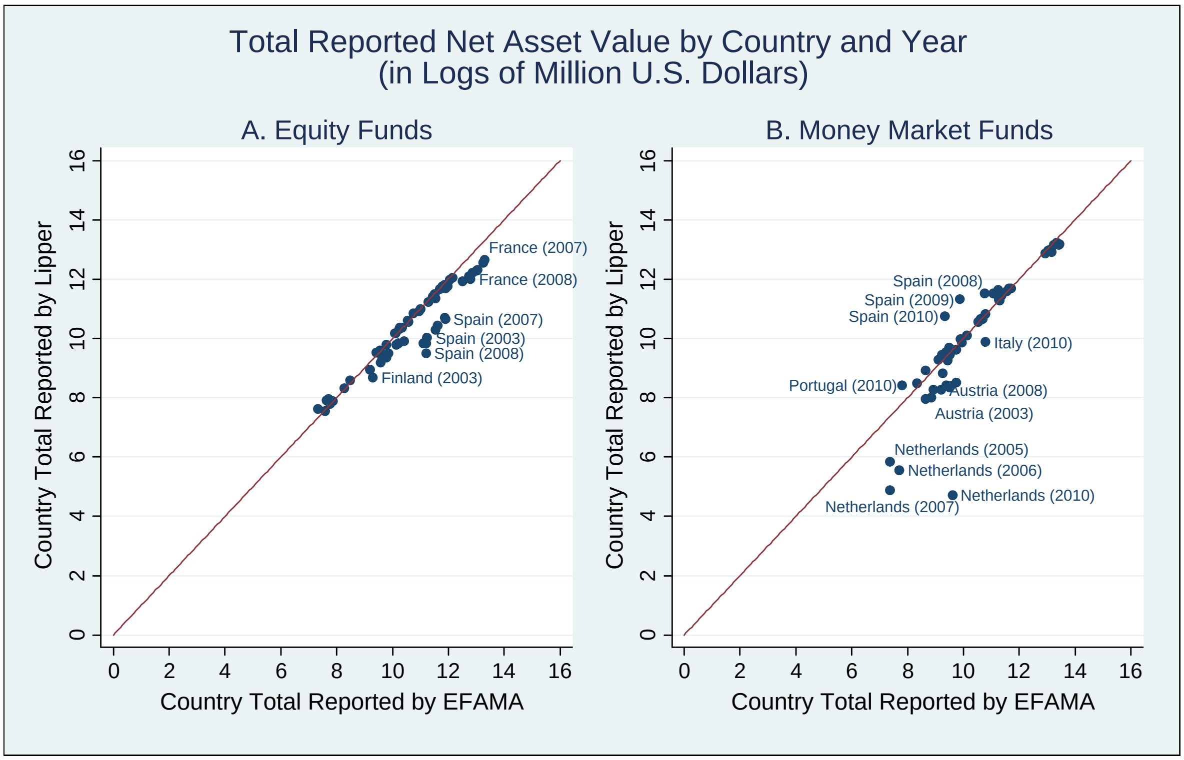

Figure 2: Plotted is the total net asset value (in the natural logarithm of million U.S. dollars) reported by the Lipper fund database on the y-axis against that reported by the European Fund and Asset Management Association (EFAMA) on the x-axis for the eight eurozone countries—Austria, Finland, France,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

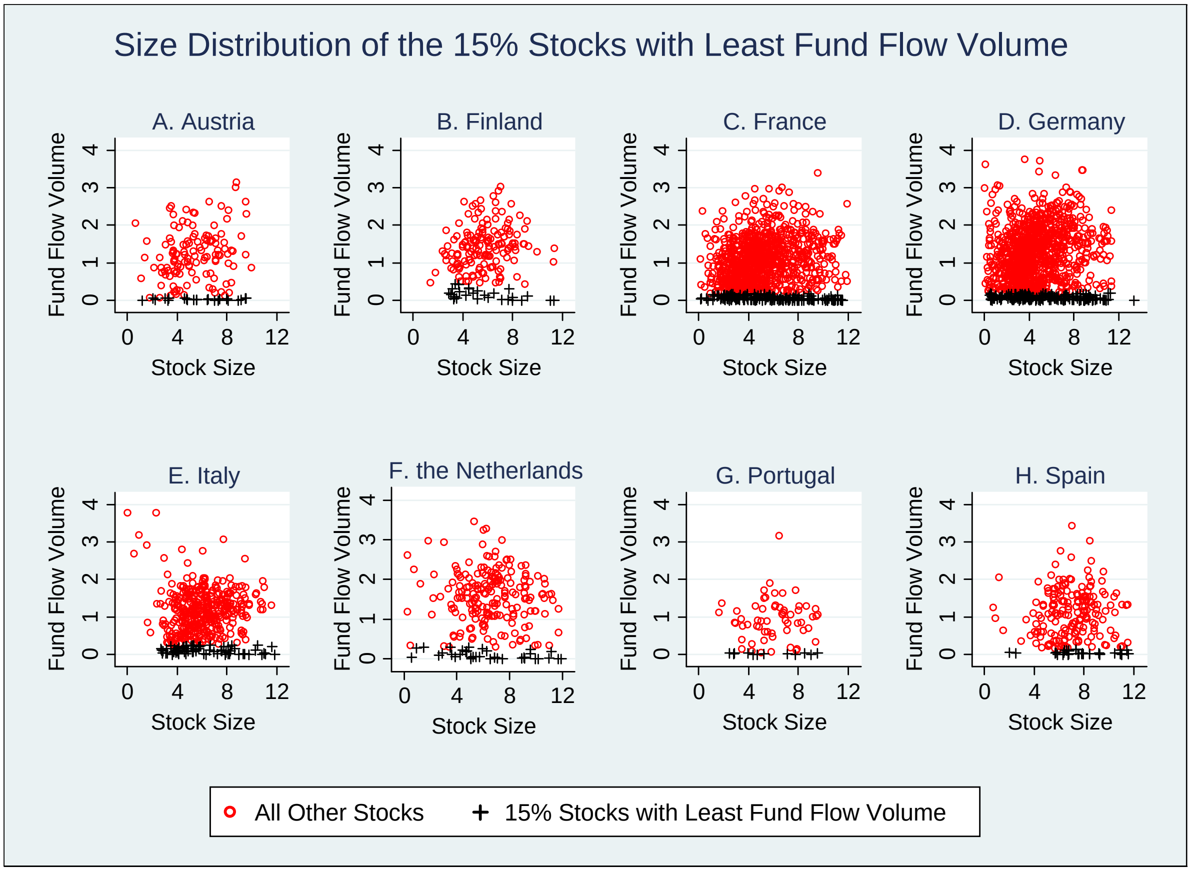

Figure 3: Plotted is the fund flow volume for stocks in eight eurozone countries against the stock size. The 15% of stocks with the lowest fund flow volume in each country are marked by black crosses, whereas all other stocks are marked by red circles. We calculate the fund flow volume for each stock as the natural logarithm of the aggregate dollar trading volume by all domestic equity funds relative to the stock's market capitalization value at the beginning of the period plus one, averaged over the sample period from 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 one, averaged over the sample period.

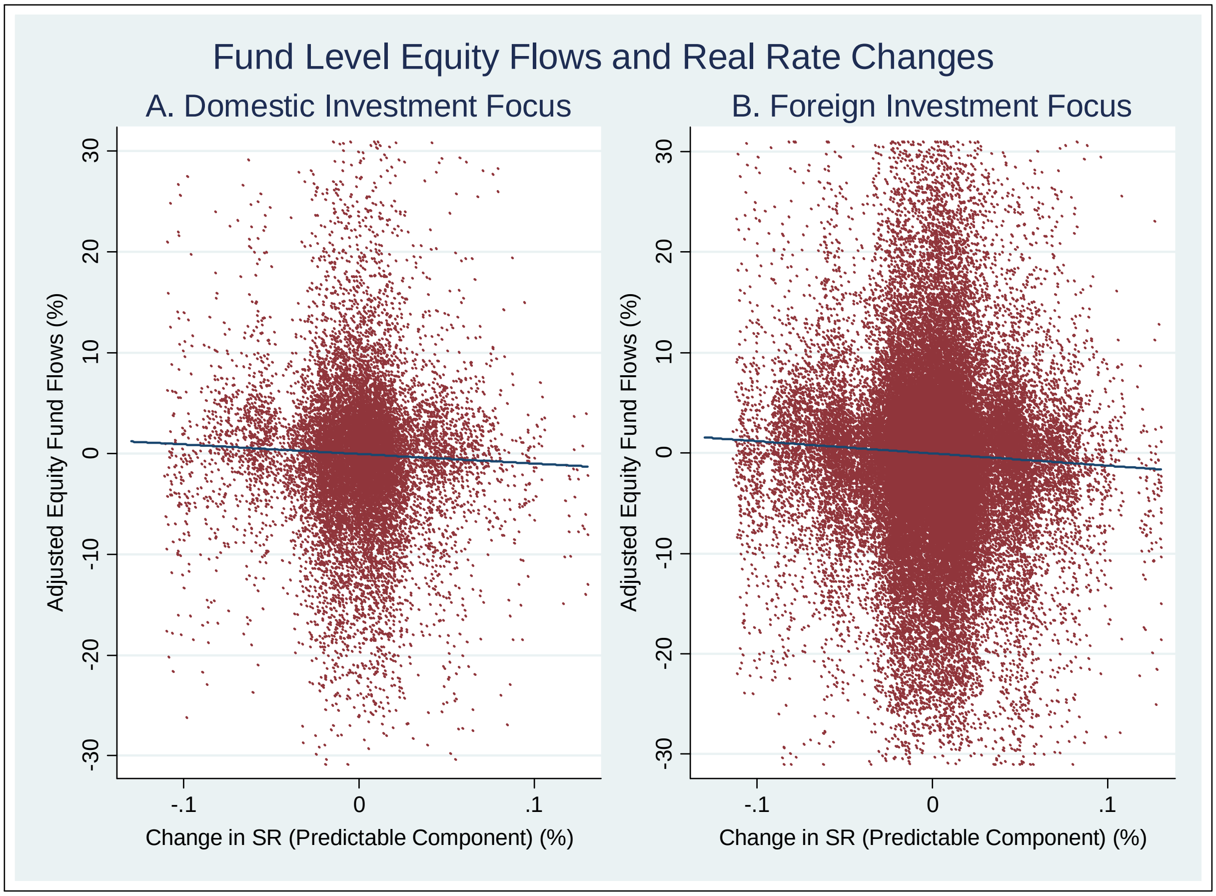

Figure 4: The figure shows the quarterly adjusted equity fund flows (from 2003/1 to 2010/4) for eight eurozone member countries against the quarterly (predicted) change of their respective real short-term interest rates (SR). Panel A plots the fund flows for equity funds with a domestic investment focus and Panel B for equity funds with a foreign investment focus. The adjusted equity fund flows denote equity fund flows minus the predictable component of fund flows from lagged fund flows ($FundFlow(-1)$ and $FundFlow(-2)$), lagged market return ($MKT(-1)$), and lagged fund returns ($FundReturn(-1)$) as estimated in Table 5, Columns 6–7. Fund flows are plotted on the y-axis and expressed in percent. On the x-axis, we plot the predictable component of the real short rate changes (in percent) of the country the fund is from; the predictable component is the one spanned by the instrument set used in Table 5, Column 6–7.

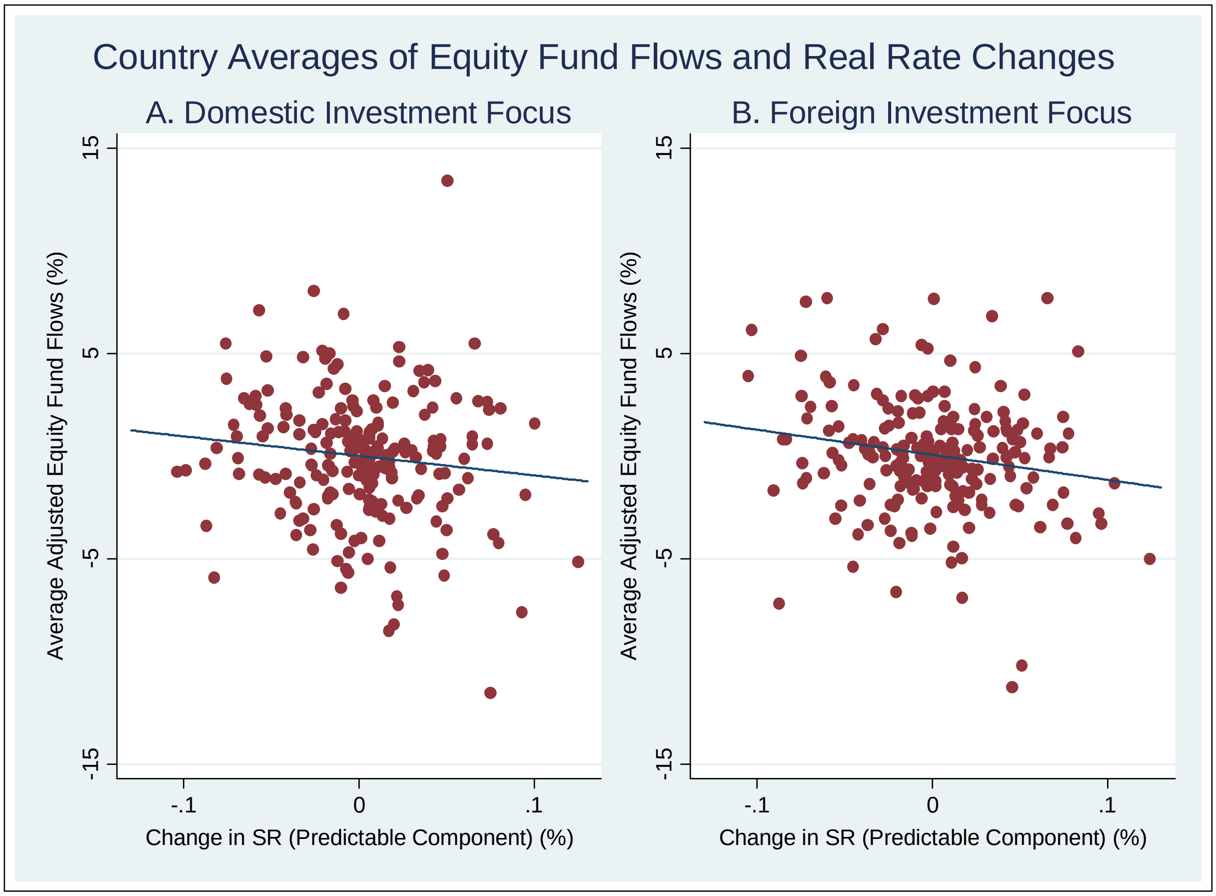

Figure 5: The figure shows the quarterly average adjusted equity fund flows (from 2003/1 to 2010/4) at the country level for eight eurozone member countries against the quarterly (predicted) change of their respective real short-term interest rates (SR). Panel A plots the average fund flow for equity funds with a domestic investment focus and Panel B for equity funds with a foreign investment focus. The adjusted equity fund flows denote equity fund flows minus the predictable component of fund flows from lagged fund flows ($FundFlow(-1)$ and $FundFlow(-2)$), lagged market return ($MKT(-1)$), and lagged fund returns ($FundReturn(-1)$) as estimated in Table 5, Columns 6–7. Fund flows are plotted on the y-axis and expressed in percent. On the x-axis, we plot the predictable component of the real short rate changes (in percent) of the country for which the average fund flows are measured; the predictable component is the one spanned by the instrument set used in Table 5, Columns 6–7.