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. A ten basis point lower real short-term interest rate is associated with a 0.7% incremental money market outflow and a 1% incremental equity market inflow by local investors relative to asset under management. The latter produces the strongest equity price increase in countries where domestic institutional investors represent 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; Gambacorta, 2009; Taylor, 2009; De Nicolo, Dell'Ariccia, Laeven, and Valenica, 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 the aggregate risk shifting as a function of the local monetary policy conditions.

Generally, monetary policy reacts to changing business conditions, which are simultaneously reflected in equity prices due to change in investor expectation. This implies that investors' reactions to monetary policy (and the subsequent stock price effect through asset reallocation) are hard to disentangle from their expectation about the stock market performance. Yet, in a currency union the central bank sets only one single short-term nominal interest rate for the entire currency area. Cross-country differences of either the short-term real 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.' Our identification strategy is similar to that of Maddaloni and Peydró (2011), who use the same cross-sectional eurozone country variations to study the effect of monetary policy on banks' risk taking. We measure cross-sectional differences in eurozone monetary conditions based on both the local short-term real interest rate and the country-specific Taylor rule residual. As explained in Appendix A, the Taylor rule residuals (TR) are retrieved 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. Alternatively, we use the local real interest rate (SR) defined as the difference between the EONIA rate and the local inflation rate to measure local monetary policy conditions.

Panel data on equity and money market flows allow us to explore the relation between monetary conditions and fund flows at both the fund level and the aggregate country level. Both the fund level and the country level panel regressions show that loose monetary policy conditions measured by the decrease in either real interest rates or Taylor rule residuals correlate strongly with the cross-sectional differences in equity fund inflows and money market fund outflows. A decrease of ten basis points in the real short-term interest rate (Taylor rule residual) is associated with a 1% (1.4%) incremental equity fund inflow relative to fund assets and a 0.7% (1.1%) incremental outflow 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. The evidence supports a powerful risk-shifting channel whereby investors react to low real rates by risk shifting from fixed-income to equity investments.

One may argue that the evidence of a strong correlation between local real interest rates and equity fund inflows has two alternative interpretations: (i) Low local real interest rates may push investors into equity fund investments. (ii) Increase in corporate profitability may pull investor flows into equity funds and simultaneously increase local inflation and reduce the local real short rate. To distinguish these two different channels, we first augment the fund flow regressions by controlling for contemporaneous change in aggregate corporate profitability (measured by the aggregate change of return on assets of locally listed stocks). We find that local cash flow shocks are not a significant determinant of equity fund inflows in our sample. Second, our data allow us to identify funds with an international investment focus, which invest more than 50% of their fund assets in foreign stocks. For those funds, any pull factor emanating from the cash flow shocks of international stocks is unlikely to correlate with the inflation rate and real short rate in the funds' domicile. Yet, our result shows that the correlation between fund flows and local real rates is similarly strong for the subsample of internationally invested funds, providing support for the argument that low local real rates push investors into equity fund investment.

Another empirical question concerns the asset price effect associated with these equity flows. Accommodating local monetary policy conditions may inflate local equity prices though (i) a lower risk-free 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 returns 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 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.

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.

2 Related Literature and Policy Issues

The role of asset prices for monetary policy is the 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.[1] Yet, such evidence matters not only for the future design of monetary policy but also for gauging the extent to which monetary policy should account for the observed asset price inflation. The current study provides direct empirical evidence on this issue in a unique currency union setting.

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

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

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 reallocation from fixed-income to 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). Therefore, it is plausible that investor risk shifting in response to monetary policy might have economically significant asset price effects beyond the direct discount rate channel. Previous works by Thorbecke (1997), Rigobon and Sack (2004), and Bernanke and Kuttner (2005) all document that expansionary (contractionary) monetary policy affects stock prices positively (negatively). Our particular contribution in relation to this strand of literature is twofold: First, based on fund flow data and its relation with local monetary policy conditions, we provide a powerful identification of how monetary policy influences investors' risky asset allocation. In an open economy, such equity fund flows provide a better measure of investor risk taking than asset prices, which are subject to many other influences.[2] Second, using the relation between local monetary policy conditions and fund flows, we can infer the stock price effect of monetary policy in a constrained structural estimation. In particular, we focus on the asset price effect of changes in the local real short rate that operate through equity market flows. Joint estimation of these flows and equity returns (relative to a local benchmark index of flow-insensitive stocks) provides a more robust inference on the asset price effect of monetary policy.

Methodologically, our study benefits from recent advances in the analysis of dynamic panels

[2] In a closed economy, net aggregate investment reallocation toward high-risk assets is not feasible; a decreased investor risk aversion implies only an asset price effect. Yet, in an open economy, reallocation to the local equity fund investment (from the less risky money market investment) can occur simultaneously with higher asset prices for those stocks that local funds hold most and are likely to channel further investment into.

(Roodman, 2006). We measure local investor risk taking based on net equity fund inflows 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 GMM (DGMM) and system GMM (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 endogenity of monetary policy, (ii) identification of investor risk-taking behavior, and (iii) quantification of the asset price effect from the enhanced risk taking by investors.

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 euro area. Within the euro area, 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 GDP growth and inflation rate; 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 euro area 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, which are suited for a causal analysis on investor behavior.

Risk shifting by local fund investors can be inferred directly from flows into those funds that are distributed and marketed exclusively in the local market given the well documented home bias in the population of fund investors (Coval and Moskowitz, 1999; Sercu and Vanpee, 2007). More risk taking amounts to outflows from locally available money market funds and simultaneous inflows into local equity funds. Such direct flow evidence provides a more solid inference on the risk-taking behavior of a large investor segment compared to indirect evidence from asset prices. Foreign investors and other domestic nonfund investors become the counterparty in this clearly defined asset reallocation problem. Unfortunately, we do not have asset allocation data for domestic nonfund investors and conjecture that they are unlikely to reverse the risk shifting of fund investors. More plausibly, the risk taking of other retail investors investing without fund intermediation might mirror the behavior of fund investors. Our empirical analysis on the asset allocation effect of monetary policy focuses on aggregate and disaggregate equity and money market fund flows and how they relate to changes in the local monetary policy conditions.

Finally, we seek to identify the linkage between monetary policy conditions and asset price inflation as well as quantify the asset price effect of enhanced risk taking 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, 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" via substitution of low yield with high yield assets. In an open economy, local fund flows from the money market to the equity market directly measure such asset substitution. 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 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 inflows 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.[3] 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

As discussed in the previous section, we can generally associate the local investor behavior with inflows and outflows of locally distributed funds because of a strong home bias in the population of fund investors. 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 (SR) or the so-called Taylor rule residuals (TR), which are the residuals obtained from a regression of the short-term nominal interest rate on both the GDP

[3]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.

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 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 euro area.

Our fund flow data are from the Lipper fund database. Fund coverage in Lipper is relatively incomplete prior to 2003. For example, it accounts for only 1.2%, 2%, and 3.3% of the entire mutual fund universe in Austria, France, and Germany, respectively, in 2002 but increases substantially to 60.3%, 68.4%, and 95.7% by the end of 2003.[4] Most funds report returns monthly, but some funds report their total net asset values only quarterly, especially in the earlier part of our sample period. Therefore, we focus our analysis on the quarterly data from January 2003 onward. 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 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 calculate

[4] The size of mutual fund industries in the euro area 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.

a fund's net quarterly flow as its net dollar flow scaled by the beginning-of-period total net asset value (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 ($FundReturn$). We then calculate the aggregate equity (money market) fund flow as the aggregate net dollar flow for all equity (money market) funds in a country scaled by these funds' aggregate beginning-of-period $TNAs$. Table 2 reports fund summary statistics.[5] Across the eurozone, investors generally withdraw capital from money market funds during our sample period. Germany and Portugal experience the largest outflows, with a mean (median) of −4.8% (−4.0%) and −3.4% (−3.3%), respectively, per quarter. By contrast, investors direct capital into equity funds in Austria, Finland, and Portugal. Across all fund-quarters, the mean (median) flow is 0.8% (−1.1%) for equity funds and −1.5% (−2.7%) for money market funds. The former register an average quarterly return of 2.2% 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 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. Benchmark stocks with extremely low fund flows exist for a wide range of stock size. The pooled mean (median) return of 3.6% (2.9%) for the $LIFI$ index (reported in Table 2) 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.

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

4 Evidence

4.1 Asset Allocation Effects of 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 market returns ($MKT_{c,t}$) 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:

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

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. But 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. Another specification concern is the endogeneity of contemporaneous covariates, $\Delta SR_{c,t}$ and $MKT_{c,t}$. Limited market depth implies that aggregate equity fund inflows could drive contemporaneous quarterly market returns so that the instrumentation of $MKT_{c,t}$ is required. Similarly, quarterly variations in factors such as country growth not only can influence the monetary policy rate (Taylor rule residuals) but also may correlate with fund flows. Such contemporaneous linkages can be purged from the estimates if lagged variables are used as instruments.

A regression based on the DGMM estimator allows for unbiased estimates with the lagged dependent variable, as well as for the instrumentation of contemporaneous covariates. Unlike LSDV, DGMM removes country fixed effects from the data through differencing. Table 3, Columns 2 and 3, report the DGMM regression results using six and nine instruments, respectively. For $\Delta SR_{c,t}$ and $MKT_{c,t}$, 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 $FundFlow_{c,t}$ 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 yield 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 short-term real interest rate by one percentage point predicts a quarterly equity fund inflow of about 10% of fund assets and a permanent inflow of about 14% (estimated by $\alpha_1/(1-\alpha_2)$). The standard deviation of quarterly changes in Taylor rule residuals is at 0.089 (reported in Table 1), which is approximately 24% smaller than the standard deviation of changes in short-term interest rates. Accordingly, we find that a decrease in the Taylor rule residual by one percentage point generates a quarterly equity inflow of about 14% of fund assets and permanent inflows of about

20%.[6] These flow effects of monetary policy are statistically highly significant and economically large. Quarterly aggregate stock market returns, $MKT_{c,t}$, do not appear to cause equity fund inflows.

An alternative estimation procedure uses 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 can only be achieved if additional orthogonality conditions for country fixed effects are met (Roodman, 2006). To be conservative, we focus on the DGMM estimates, but report the SGMM results as a robustness check.[7] 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.[8]

Table 4 provides the corresponding results for money market flows. The estimated autocorrelation for money market flows is between 0.31 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 8.2, 7.7, and 7.1 for LSDV, DGMM1, and DGMM2, respectively. These results are all statistically significant at the 5% level or better. Using Taylor rule residuals instead of short-term real interest rates in Panel B again shows that the estimated flow effects are large: A loose monetary policy with the Taylor rule residual lowered by one percentage point generates an immediate incremental money market outflow of approximately 11% of fund assets and a permanent effect of roughly 15.7% ($\approx 11\%/(1-0.3)$). The SGMM estimates in Columns 4 and 5 are similar to those of DGMM. The validity of identification restrictions is not rejected, even under SGMM2, in which 12 instruments are used.

Overall, the aggregate flow regressions show a quantitatively strong risk shifting into equity

[6] The standard deviation of changes in short-term real interest rates is only 0.117 percentage point in our sample. But monetary policy adjustments are often undertaken over several quarters, and a one-percentage-point change in the real rate (possibly implemented over a few quarters) does not represent an implausibly large policy change.

[7] 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.

[8] 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.

fund investment in a loose monetary policy environment. 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 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 contemporaneous fund performance ($FundReturn_{j,t}$) and lagged fund performance ($FundRetrun_{j,t-1}$ and $FundRetrun_{j,t-2}$):

$$\begin{aligned} FundFlow_{j,t} &= \alpha_0 + \alpha_1 \Delta SR_{c,t} + \alpha_2 FundFlow_{j,t-1} + \alpha_3 FundFlow_{j,t-2} + \\ &\quad + \alpha_4 MKT_{c,t} + \alpha_5 FundReturn_{j,t} + \alpha_6 FundRetrun_{j,t-1} + \mu_j + \epsilon_{j,t}. \end{aligned} \tag{2}$$

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. 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 4.3.

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 specifications in Table 5, Columns 1 and 2, are 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 results in Panel A, Columns 1 and 2 (with and without fund performance control) to the corresponding DGMM results in Columns 3 and 4 shows that the former yields an estimated autocorrelation of 0.19 for fund flows, which is only slightly more than half 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. Aggregate market returns, MKT, again have no reliable explanatory power in the DGMM regressions, consistent with the findings from Tables 3 and 4. By contrast, contemporaneous and lagged fund returns are highly significant determinants of equity flows. The more eleborate specification labeled DGMM2 in Table 5 implies that a 1% higher quarterly fund return correlates with a short-run (contemporaneous) inflow of about 0.3% of asset values and a lagged effect of roughly 0.13%.

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 Hansen test does not reject that all (over-)identifying restrictions are simultaneously fulfilled. The equity flow results are also robust to the alternative specification of system GMM, reported in Column 5.

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 11 (based on the estimates in DGMM reported in Columns 3 and 4), compared to the corresponding estimate of about 13 for the SGMM reported in Column 5. The coefficient estimtates for ΔSR are all statisitcally 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 inflows 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 output 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.

What is the scope for a *direct macroeconomic channel* on investor flows under the observed negative correlation between equity fund flows and change in the real short rate? 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 lower corporate profitability. Here, positive equity fund inflows would occur parallel to higher inflation only if local investors are contrarian equity investors.

In either case, direct local investor reaction to local aggregate output or firm profitability 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 and profitability measures a better proxy for contemporaneous macroeconomic shocks than the measure of real short rates.

In Table 5, Column 4, we augment the baseline regression by the quarterly changes in local firm profitability, measured by the aggregate return on assets (ΔROA) of locally listed stocks, and the national GDP growth ($gGDP$). The result reported in Table 5, Column 6 for equity funds shows that neither of the two control variables attenuates the correlation coefficient between changes in the real short rate and the net equity fund inflows. In particular, the two variables ΔROA and $gGDP$ are both statistically insignificant, and the point estimate of ΔSR, -9.471 (t-stat $= -5.10$), is quantitatively similar to the estimate of -9.889 (t-stat $= -5.34$) for the baseline regression reported in Column 4. In Table 6, Column 6, we report the augmented regression result for money market funds. Again, we find that the point estimate of ΔSR is nearly unchanged with the inclusion of the two additional variables ΔROA and $gGDP$. Hence, we find no evidence for a direct macro channel on local investor flows that may lead us to attribute a non-causal role to the real short rate.

A further test of the causal effect consists in focusing on the equity flows into those funds with more than 50% of their assets invested in foreign stocks. Profitability shocks to foreign stocks are less likely to feature any meaningful correlation with the inflation rate or the real short rate of the funds' domicile, thereby reducing the scope for causal effects from firm level shocks to changes in country specific real short rates and local investors' equity inflows. The flow regression reported in Table 5, Column 7, is exclusively for funds with a foreign stock investment focus, with a sample size of $58,300$ observations compared to the full sample of $73,767$ observations. We find a similarly strong correlation between fund flows and local real rates for this subsample of funds. The point estimate of ΔSR is -11.381 (t-stat $= -5.21$), compared to the estimate of -9.889 (t-stat $= -5.34$) for the full sample.

We conclude that the equity flows are not caused by firm level profitability shocks to listed stocks that simultaneously influence the local real short-term interest rate and fund inflows. Instead, the strong correlation between equity fund inflows and lower local real rates are likely to reflect investor risk shifting from fixed-income to equity investment under loose monetary policy conditions captured by the real short rate. The following section seeks to isolate and quantify the asset pricing effect of such risk shifting.

4.3 Stock Price Effects of Monetary Policy

Previous empirical research shows that aggregate fund flows relate to sizeable stock price effects (Jotikasthira, Lundblad, and Ramadorai, 2012). Section 4.2.1 seeks to quantify this relationship for equity fund flows triggered by variations in local monetary policy conditions across eurozone countries. In Section 4.2.2, we discuss whether changing local risk premia can account for the aggregate stock price dynamics; the discussion is followed by more robustness considerations in Section 4.2.3.

4.3.1 Simultaneous Equation Approach

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. 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 inflows 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 3-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 pressure as the fund return in excess of the benchmark return:

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

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

The second identifying step consists in 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 due to 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 return and fund return 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 inflows.

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 in order 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 (reported to the Factset institutional ownership database) varies from 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 $LocInstShare(c)$, which measure local institutional investment relative to the size of the local stock market. This puts more weight on fund flows in locations where 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 3 and 4. 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 -9.3 (in DGMM1). 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. Inversely, if the home bias is small, an accomodating monetary policy exports 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.

4.3.2 Price Pressure vs. Risk Premium Changes

The flow-related price difference between the stocks in the fund investment universe and the stocks in the benchmark *Low Investor Flow Index* ($LIFI$) has two possible interpretations. First, the flow-related price pressure may generate aggregate stock price effects. This interpretation presupposes a certain degree of market segmentation between fund stocks and the benchmark stocks as well as limits to arbitrage. The relative illiquidity of the benchmark $LIFI$ stocks may impair effective arbitrage between the two stock groups. Moreover, the aggregate nature of the mispricing also would require considerable arbitrage capital to eliminate mispricing. Second, fund flows may just occur contemporaneously with changes in local risk premia, coupled with differential risk exposures of fund stocks and benchmark stocks to these changing risk premia. In this second case, stocks are priced (to some degree) locally so that changing risk aversion of local investors changes the local market risk premium (and/or other risk premia).[9] Unlike changes in *local risk premia*, changes in *global risk premia* should have no explanatory power over the observed price inflation of the flow-sensitive stocks because they cannot be aligned with changes in local monetary conditions measured by the cross-sectional deviations from the eurozone average.

Empirically, it is difficult to gauge the relative accountability of the two channels (a simple price pressure channel and a conditional asset pricing explanation) for the observed asset price effect documented in Table 7 because the time variation of investor risk aversion and market risk premia is not directly observable. Nonetheless, in Table 8 we explore whether local betas differ between the 15% of stocks included in the benchmark $LIFI$ index and the remaining 85% of stocks. Recent studies (e.g., Bekaert, Hodrick, and Zhang, 2009; Eun et al., 2010; Hou, Karolyi, and Kho, 2011; and Karolyi and Wu, 2012) emphasize the importance of both local and global components of international stock returns. We therefore employ an international 8-factor model comprising four domestic factors (i.e., market, SMB, HML, and MOM factors) and four corresponding international factors to estimate the betas for each individual stock, using weekly data over the whole sample period. Appendix B describes in detail the construction of the weekly domestic and international factors. Because the constituent stocks of the $LIFI$ index change every six months, Table 8 reports the time-series average of the value-weighted betas for the portfolio of benchmark $LIFI$ stocks in Panel A and nonbenchmark stocks in Panel B. The average number of stocks included in each respective portfolio is reported in Column 1.

In all eight sample countries, the benchmark $LIFI$ index stocks feature lower domestic market betas and generally lower domestic risk loadings than the complementary set of (non-benchmark) stocks in which funds invest most. For the pooled sample, we find that fund stocks have a local market beta of 0.917 compared to 0.646 for the benchmark stocks. The difference

[9] For recent evidence on local versus global risk pricing of stocks, see Hau (2011).

in the market beta of 0.271 implies that changes in the local market risk premium can partially explain the differences in the return dynamics between fund stocks and benchmark stocks. For example, a decrease of the local market premium from 6% to 4% and a risk-less rate of 4% imply that the return of fund stocks should exceed that of the benchmark stocks by roughly 4%.

Ultimately, without precise information on the magnitude of change in local investor risk aversion and market risk premia in relation to local monetary policy conditions, this paper remains agnostic about the dominant force of the asset price inflation identified in the previous subsection. We consider further analysis of asset pricing issues beyond the scope of this paper and leave it to future research.

4.4 Robustness

Disaggregate fund flow regressions (in Tables 5 and 6) discussed earlier use short-term real interest rates as the measure for local monetary policy conditions. In our first robustness test, we repeat these disaggregate fund flow regressions by replacing changes in real short rates, ΔSR, with the corresponding changes in Taylor rule residuals, ΔTR. The results (reported in Tables A1 and A2 of the Web Appendix) are qualitatively very similar to those reported in Tables 5 and 6. For example, the point estimates for ΔTR are -14.293 (t-stat= -5.07) and 15.138 (t-stat=2.12) for equity funds and money market funds, respectively, compared to the corresponding estimates of -9.889 (t-stat=-5.34) and 10.930 (t-stat=2.24) for ΔSR in DGMM2. The numerically slightly larger point estimates for the ΔTR coefficient reflects the fact that the standard deviation of the change in Taylor rule residuals is about 24% smaller than the corresponding standard deviation for the change in the real short rate. The disaggregate fund flow results are therefore robust to the two alternative measures of the monetary policy rate.

The second 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 (reported in Tables A3 and A4 of the Web Appendix) again produce very similar point estimates for the effect of changes in the real short rate on equity and money market flows. Take DGMM2 estimates for example. The point estimate for ΔSR is -9.091

(t-stat$= -5.98$) for equity funds and 10.282 (t-stat$= 2.26$) for money market funds under the equal-weighted approach, compared to -9.889 (t-stat$= -5.34$) and 10.930 (t-stat$= 2.24$) under the value-weighted appraoch. We conclude that the interest rate effect on fund flows does not depend on fund size.

The third robustness test concerns the alternative threshold for constructing the Low Investor Flow Index ($LIFI$) index. 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. As a robustness check, we use an alternative threshold of 10% or 20%. The results are reported in Tables A8 and A9 of the Web Appendix. The total return effect of ΔSR is approximately -39 (-15) and -22 (-11) for the 10% and 20% thresholds, respectively, using $LocInstShare(c)$-adjusted (equal) country weights, compared to the return effect of -34 (-14) for the 15% threshold. Overall, the quanitative 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.

Lastly, we consider 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 then repeat the simultaneous equation regressions of Table 7 using this alternative index as the relevant return benchmark. We find similar results. Specifically, 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 $\widehat{\beta}_1 + \widehat{\beta}_2 + \widehat{\beta}_3 \approx -36$. These results are provided in Table A7 of the Web Appendix. 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.

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 features a tight link between the risk-taking decisions of retail investors and fund flows to equity and money market funds in the respective countries.

First, we find that loose local monetary policy conditions (measured by 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 asset reallocation 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. On average, a one-percentage-point lower short-term real interest rate is associated with an 7% incremental money market outflow and a 10% incremental equity market inflow relative to fund assets under management.

Second, 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 the local sentiment 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.

We interpret our evidence as support for a powerful link between monetary policy and investors' asset allocation decisions. Loose monetary policy appears to diminish investor risk aversion 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 on investors' asset allocation decisions can serve as a useful complementary source of information on investor risk choices.

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
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 $\delta_2 = 0.658$ [$t = 11.78$]. There total number of observations is 256, and the adjusted R-squared is 0.349.	Datastream
MKT	Quarterly return on the MSCI country market index.	Datastream
$LIFI$	Quarterly return on the value-weighted index of the 15% local stocks with the lowest absolute fund in- 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

Appendix A continued.

Variable	Description	Source
$LFHI$	Quarterly return on the value-weighted index for 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 $TNAs$.	Lipper

Appendix B. Risk Factor Construction

Risk factors are constructed are based on weekly stock returns in U.S. dollars from Datastream over the eight-year period from July 2002 to December 2010. We exclude non-common stocks such as REITs, closed-end funds, warrants, etc. We also exclude firms that are incorporated outside their home countries, as well as those indicated by Datastream as duplicates. To filter out the recording errors in Datastream, we assign missing values to R_t and R_{t-1} if $(1+R_t)(1+R_{t-1}) < 0.5$ and at least one of them is greater than or equal to 200%. R_t is the stock return in month t. In addition, in view of Datastream's practice to set the return index to a constant once a stock ceases trading, we treat those constant values as missing values in the inactive file.

In the first step, we determine domestic factors for each country. The domestic market factor is given by the excess return in U.S. dollars of the country's equity index return over the U.S. Treasury Bill rate. We calculate country index returns using the MSCI country market indices obtained from Datastream. For the size and book-to-market factors we follow a methodology similar to Fama and French (1993). All stocks reporting a market capitalization as of the last Friday of June and a positive book-to-market ratio at the end of the previous year are double sorted into two size groups and three book-to-market classifications. Half of the stocks are classified as large-cap (B) and the other half as small-cap (S). For the book-to-market classification, the bottom 30% of firms are classified as L, the middle 40% as M, and the highest 30% as H. The intersection of the rankings allows for six value-weighted portfolios: HB, MB, LB, HS, MS, and LS. Formally, we define

$$SMB = \frac{1}{3}(HS + MS + LS) - \frac{1}{3}(HB + MB + LB)$$
$$HML = \frac{1}{2}(HB + HS) - \frac{1}{2}(LB + LS).$$

The weekly Friday-to-Friday returns for SMB and HML are then calculated from the first week of July in one year to the last week of June in the next. The momentum factor (MOM) is re-balanced every month; we rank stocks at the last Friday of month $t-1$ based on their cumulative returns from months $t-13$ to $t-2$ (i.e., prior 2–12 month returns by skipping month $t-1$) and market value at the last Friday of month $t-1$. Stock inclusion in the portfolio construction requires nonmissing values for the cumulative return and market value. For the market-cap classification, half of the stocks are again classified as large-cap (B) and the other half as small-cap (S). For the past returns classification, the bottom 30% are classified as LR (low return), the middle 40% as MR (middle return), and the highest 30% as HR (high return). The momentum factor is defined as

$$MOM = \frac{1}{2}(SHR + BHR) - \frac{1}{2}(SLR + BLR).$$

The weekly MOM returns are then calculated from the first week to the last week of month t. For the U.S. factors, we use the data posted on Kenneth R. French's website. If a country has fewer than 100 stocks qualifying for the portfolio construction, we set SMB, HML, and MOM factors as missing for the respective year.

A country's international factors are calculated in a second step as the weighted average of the respective domestic factors of all other countries. The weights are given by the relative stock market capitalization of each foreign country at the beginning of the year. The stock market capitalization data is obtained from the World Development Indicator. Equal-weighted regressions in Panel B give qualitatively similar fund return results, suggesting that the flow-performance relation for European equity funds is not very heterogenous across fund size.

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Table 1: Summary Statistics of Macroeconomic Variables

Reported are the summary statistics of the average quarterly overnight interest rates for the euro area ($EONIA$) and the average quarterly real GDP growth ($gGDP$), inflation rates (INF), and aggregate change in return on assets (ΔROA) for the sample countries. The sample consists of Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain during the period from 2003/1–2010/4. We also report the short-term real interest rates (SR) and the Taylor rule residuals (TR) by country as well as their cross-country averages. The cross-country averages for change in short-term real interest rates (ΔSR) and change in Taylor rule residuals (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 ×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
Short-Term Real Interest Rate (SR) ×100						
Austria	32	0.101	0.118	0.246	−0.399	0.506
Finland	32	0.220	0.308	0.239	−0.500	0.548
France	32	0.140	0.126	0.250	−0.312	0.678
Germany	32	0.182	0.193	0.192	−0.221	0.501
Italy	32	0.053	0.031	0.224	−0.293	0.594
Netherlands	32	0.165	0.145	0.259	−0.274	0.672
Portugal	32	0.049	−0.014	0.268	−0.440	0.468
Spain	32	−0.096	−0.155	0.260	−0.480	0.408
All SR	256	0.102	0.101	0.258	−0.500	0.678
All ΔSR	256	−0.016	−0.008	0.117	−0.411	0.333
Taylor Rule Residual (TR) ×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 fund flows measured over 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 in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 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 cross-sectional means. Column 1 provides the estimate using the LSDV regression. Columns 2 and 3 and 4 and 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 with the monetary policy rate proxied by the short-term real interest rate. Changes (from the previous quarter) in the short-term real interest rates and the Taylor rule residuals are denoted by ΔST and ΔTR, respectively; $FundFlow(-1)$ denotes the fund flow in the previous quarter; MKT is the contemporaneous country stock market return. 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	−4.361	−9.556	−9.675	−9.592	−10.042
	[−2.02]	[−4.07]	[−4.34]	[−4.62]	[−5.34]
$FundFlow(-1)$	0.348	0.312	0.339	0.219	0.287
	[4.36]	[3.04]	[3.50]	[1.65]	[2.34]
MKT	0.076	0.076	0.072	0.062	0.058
	[2.29]	[1.15]	[1.09]	[1.06]	[0.98]
Obs.	254	246	246	254	254
$Adj.R^2$	0.305				
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 1-2	Lags 1-2	Lags 1-2	Lags 1-2
Total Number		6	9	9	12
$AR(1)$		0.012	0.009	0.026	0.015
$AR(2)$		0.484	0.426	0.677	0.515
Hansen Test		0.393	0.515	0.328	0.735

Panel B: Taylor Rule Residuals

Dep. Variable: Fund Flow	LSDV (6)	DGMM1 (7)	DGMM2 (8)	SGMM1 (9)	SGMM2 (10)
ΔTR	−6.032	−13.969	−14.166	−14.703	−15.484
	[−1.82]	[−3.98]	[−4.20]	[−4.37]	[−5.02]
$FundFlow(-1)$	0.368	0.306	0.328	0.263	0.326
	[4.32]	[2.76]	[3.00]	[2.41]	[3.15]
MKT	0.061	0.046	0.041	0.053	0.048
	[2.30]	[0.96]	[0.88]	[0.96]	[0.87]
Obs.	248	240	240	248	248
$Adj.R^2$	0.318				
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 1-2	Lags 1-2	Lags 1-2	Lags 1-2
Total Number		6	9	9	12
$AR(1)$		0.022	0.019	0.027	0.019
$AR(2)$		0.318	0.284	0.388	0.304
Hansen Test		0.380	0.638	0.391	0.749

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 in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 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 cross-sectional means. Column 1 provides the estimate using the least square dummy variable (LSDV) regression. Columns 2 and 3 and 4 and 5 provide the estimates using one-step 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 short-term real interest rate. Changes (from the previous quarter) in the short-term real interest rates and the Taylor rule residuals are denoted by ΔST and ΔTR, respectively; $FundFlow(-1)$ denotes the fund flow in the previous quarter; MKT is the contemporaneous country stock market return. 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	8.199	7.683	7.148	8.186	7.940
	[2.32]	[2.24]	[2.00]	[2.66]	[2.61]
$FundFlow(-1)$	0.364	0.362	0.315	0.370	0.316
	[5.04]	[5.08]	[4.66]	[6.14]	[5.24]
MKT	0.015	−0.018	−0.010	0.001	0.003
	[0.21]	[−0.17]	[−0.10]	[0.01]	[0.03]
Obs.	249	240	240	249	249
$Adj.R^2$	0.225				
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 1-2	Lags 1-2	Lags 1-2	Lags 1-2
Total Number		6	9	9	12
AR(1)		0.009	0.008	0.009	0.007
AR(2)		0.901	0.967	0.873	0.961
Hansen Test		0.411	0.379	0.692	0.747

Panel B: Taylor Rule Residuals

Dep. Variable: Fund Flow	LSDV (6)	DGMM1 (7)	DGMM2 (8)	SGMM1 (9)	SGMM2 (10)
ΔTR	12.665	12.050	11.087	11.610	11.347
	[2.45]	[2.23]	[1.96]	[2.25]	[2.17]
$FundFlow(-1)$	0.360	0.363	0.314	0.365	0.316
	[4.94]	[5.08]	[4.64]	[5.82]	[5.21]
MKT	0.038	0.001	0.008	0.008	0.011
	[0.54]	[0.01]	[0.08]	[0.11]	[0.14]
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 1-2	Lags 1-2	Lags 1-2	Lags 1-2
Total Number		6	9	9	12
$AR(1)$		0.010	0.009	0.011	0.008
$AR(2)$		0.798	0.865	0.783	0.862
Hansen Test		0.395	0.410	0.607	0.652

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 in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 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 rates ΔSR; (ii) fund flows at lags 1 and 2 given by $FundFlow(-1)$ and $FundFlow(-2)$, respectively; (iii) the country stock market return MKT; (iv) individual fund returns in the current and previous quarter given by $FundReturn$ and $FundReturn(-1)$; (v) change in aggregate corporate profitability, proxied by change in return on assets (ΔROA) at the country-level; and (vi) GDP growth ($gGDP$). Columns 1 and 2 state the results for the least square dummy variable (LSDV) regressions without instruments. Columns 3 and 4 provide the estimates using the difference generalized method of moments (DGMM) estimator, whereas Column 5 reports estimates based on the system generalized method of moments (SGMM). Column 6 uses the same setup as Column 4 but includes two additional regressors, ΔROA and $gGDP$. Column 7 provides the DGMM estimates for a subsample of funds, which invest more than 50% of their fund assets in foreign securities. All regressions report robust t-statistics in brackets. Also reported are the number of observations (*Obs.*), adjusted R-square for the LSDV regression ($Adj.R^2$), type and total number of instruments used in each specification, p-values for the tests of the first and second order autocorrelations of the residuals [$AR(1)$ and $AR(2)$], and Hansen test for the overidentification conditions. Appendix A provides the variable definitions in detail.

Dep. Variable: Fund Flow	LSDV1 (1)	LSDV2 (2)	DGMM1 (3)	DGMM2 (4)	SGMM (5)	DGMM3 (6)	DGMM4 (7)
ΔSR	−3.631	−4.553	−9.328	−9.889	−8.758	−9.471	−11.381
	[−3.51]	[−4.35]	[−5.27]	[−5.34]	[−5.59]	[−5.10]	[−5.21]
$FundFlow(-1)$	0.195	0.189	0.351	0.341	0.344	0.340	0.348
	[12.29]	[12.09]	[14.32]	[14.33]	[14.26]	[14.28]	[13.32]
$FundFlow(-2)$	0.061	0.061	0.129	0.127	0.156	0.127	0.118
	[5.72]	[5.71]	[4.83]	[4.80]	[5.63]	[4.81]	[4.14]
MKT	0.051	0.016	0.084	0.039	0.053	0.040	0.044
	[2.46]	[0.78]	[2.49]	[1.27]	[1.72]	[1.30]	[1.22]
$FundReturn$		0.193		0.301	0.239	0.308	0.366
		[9.81]		[5.85]	[5.69]	[6.08]	[5.51]
$FundReturn(-1)$		0.131		0.134	0.135	0.138	0.162
		[6.89]		[3.73]	[3.75]	[3.83]	[3.52]
ΔROA						−0.079	
						[−0.34]	
$gGDP$						0.172	
						[0.61]	
Obs.	78,735	78,735	73,767	73,767	78,735	73,767	58,300
$Adj.R^2$	0.158	0.169					
Instruments							
ΔSR			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
MKT			Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundReturn$				Lags 1-2	Lags 1-2	Lags 1-2	Lags 1-2
ΔROA						Lags 1-2	
$gGDP$						Lags 1-2	
Total Number			6	8	12	12	8
AR(1)			0.000	0.000	0.000	0.000	0.000
AR(2)			0.772	0.609	0.243	0.629	0.791
Hansen Test			0.244	0.409	0.000	0.000	0.924

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 in Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain over the period 2003/1–2010/4. Similar to the setup 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 Columns 1–6 of Table 5.

Dep. Variable: Fund Flow	LSDV1 (1)	LSDV2 (2)	DGMM1 (3)	DGMM2 (4)	SGMM (5)	DGMM3 (6)
ΔSR	12.558	11.999	11.291	10.930	12.705	10.393
	[3.41]	[3.32]	[2.19]	[2.24]	[3.68]	[2.14]
$FundFlow(-1)$	0.150	0.142	0.291	0.287	0.283	0.284
	[3.26]	[3.16]	[3.20]	[3.18]	[3.17]	[3.13]
$FundFlow(-2)$	0.004	0.003	0.103	0.105	0.103	0.103
	[0.14]	[0.09]	[1.87]	[1.90]	[1.97]	[1.90]
MKT	0.065	0.048	0.096	0.105	0.049	0.108
	[1.03]	[0.75]	[1.13]	[1.21]	[0.84]	[1.22]
$FundReturn$		0.738		1.384	0.645	1.342
		[2.12]		[2.15]	[2.35]	[2.20]
$FundReturn(-1)$		0.343		0.447	0.084	0.390
		[1.50]		[0.98]	[0.26]	[0.85]
ΔROA						0.155
						[0.35]
$gGDP$						0.760
						[0.92]
Obs.	19,694	19,694	17,659	17,659	19,694	17,659
Adj.R^2	0.113	0.123				
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 1-2	Lags 1-2	Lags 1-2	Lags 1-2
$FundReturn$				Lags 1-2	Lags 1-2	Lags 1-2
ΔROA						Lags 1-2
$gGDP$						Lags 1-2
Total Number			6	8	12	12
AR(1)			0.000	0.000	0.000	0.000
AR(2)			0.250	0.267	0.365	0.253
Hansen Test			0.559	0.380	0.215	0.615

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 rates 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 cross-sectional means. The sample covers all locally 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 aggregate local investment of all local institutional investors relative to the local stock market capitalization. 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.*), p-values for the two linear constraints on the flow dynamics, type and number of instruments, and p-value of the Hansen overidentification test 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 Coeffients	−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
Constraint 1 (p value)	0.000	0.000	0.000	0.000	0.000	0.000
Constraint 2 (p value)	0.000	0.000	0.000	0.000	0.000	0.000

Table 8: Risk Loadings for the *LIFI* Benchmark and Non-*LIFI* Portfolio

We report the time-series average of the value-weighted betas for the *LIFI* benchmark stocks (Panel A) and non-*LIFI* stocks (Panel B) with respect to an eight-factor model, comprising four domestic factors (i.e., market, SMB, HML, and MOM factors) and four corresponding international factors. The betas for each individual stock are estimated using weekly data from January 2003 to December 2010. The *LIFI* index comprises the 15% of stocks with the lowest fund flows in each country over the previous three-year period and is re-balanced every six months. The average number of stocks included in the portfolio is reported in Column 1. Appendix B describes the construction of the factors in detail. We also test whether the distribution of stock betas differs between the benchmark stocks and their complementary nonbenchmark stocks and mark any statistically significant difference (according to the Wilcoxon rank-sum test) at the 5%, 3%, and 1% significance levels by *,**, and ***, respectively.

Panel A: Risk Loading for *LIFI* Portfolio

		Domestic Betas				International Betas			
	Obs.	Market	SMB	HML	Mom.	Market	SMB	HML	Mom.
Austria	13	0.407***	0.501***	0.135***	0.062***	0.018**	0.121***	−0.276***	0.139***
Finland	20	0.681***	0.464***	0.747***	−0.068***	0.217***	−0.162***	−0.177***	−0.074***
France	95	0.812***	0.421***	−0.141***	−0.028*	0.102***	0.178***	0.031***	−0.024***
Germany	105	0.569***	0.745***	0.107***	−0.098***	−0.071***	0.063***	0.103***	0.169***
Italy	45	0.814***	1.043***	0.414***	−0.033**	0.043	−0.118***	0.041***	0.006***
Netherlands	22	0.739***	0.188***	−0.190***	0.010***	−0.043***	−0.134***	0.082**	0.142***
Portugal	8	0.509***	0.155**	0.206***	−0.222***	0.043**	0.121	−0.252***	0.224***
Spain	23	0.638***	0.401***	0.086***	−0.005	0.080	0.436***	−0.212***	0.169***
All *Stocks*	331	0.646***	0.490***	0.171***	−0.048	0.049*	0.063	−0.082***	0.094***

Panel B: Risk Loading for Non-*LIFI* Portfolio

		Domestic Betas				International Betas			
	Obs.	Market	SMB	HML	Mom.	Market	SMB	HML	Mom.
Austria	76	0.801***	0.021***	0.020***	0.042***	0.045**	0.080***	0.054***	−0.061***
Finland	110	1.019***	0.030***	0.029***	−0.048***	−0.083***	−0.007***	0.032***	−0.038***
France	556	0.966***	0.035***	0.043***	−0.030*	−0.019***	0.082***	0.062***	−0.059***
Germany	618	0.826***	0.082***	0.061***	−0.149***	0.101***	0.035***	0.176***	0.026***
Italy	262	0.924***	0.213***	0.039***	−0.047**	0.047	0.108***	−0.067***	0.051***
Netherlands	124	0.972***	−0.005***	0.022***	−0.047***	−0.030***	0.106***	0.047**	−0.116***
Portugal	43	0.903***	−0.063**	0.020***	0.019***	0.026**	0.074	0.149***	−0.094***
Spain	131	0.922***	0.138***	0.010***	−0.021	0.070	0.096***	0.040***	−0.010***
All *Stocks*	1,920	0.917***	0.056***	0.030***	−0.035	0.020*	0.072	0.061***	−0.038***

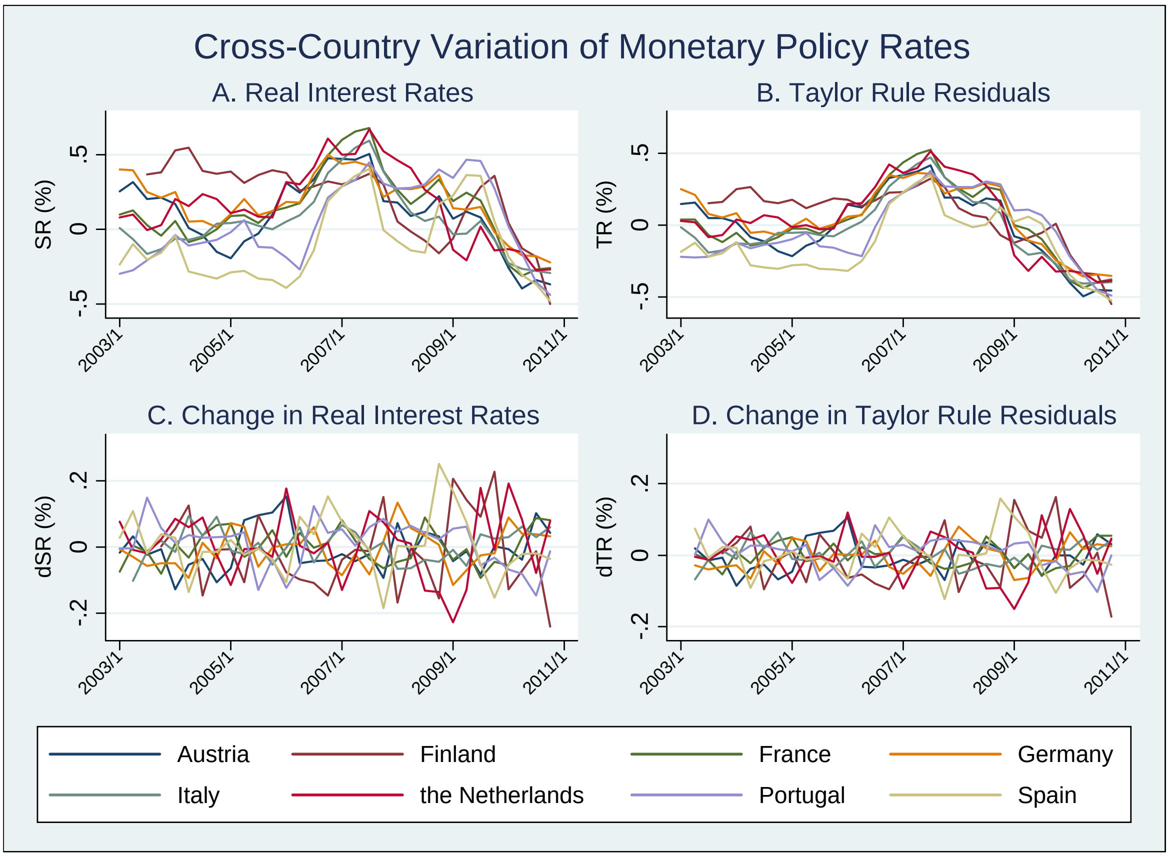

Figure 1: Plotted in Panels A and B are the short-term real interest rates (SR) and the quarterly Taylor rule residuals (TR), respectively, for each of the eight eurozone countries–Austria, Finland, France, Germany, Italy, the Netherlands, Portugal, and Spain from 2003/1–2010/4. Panels C and D plot the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

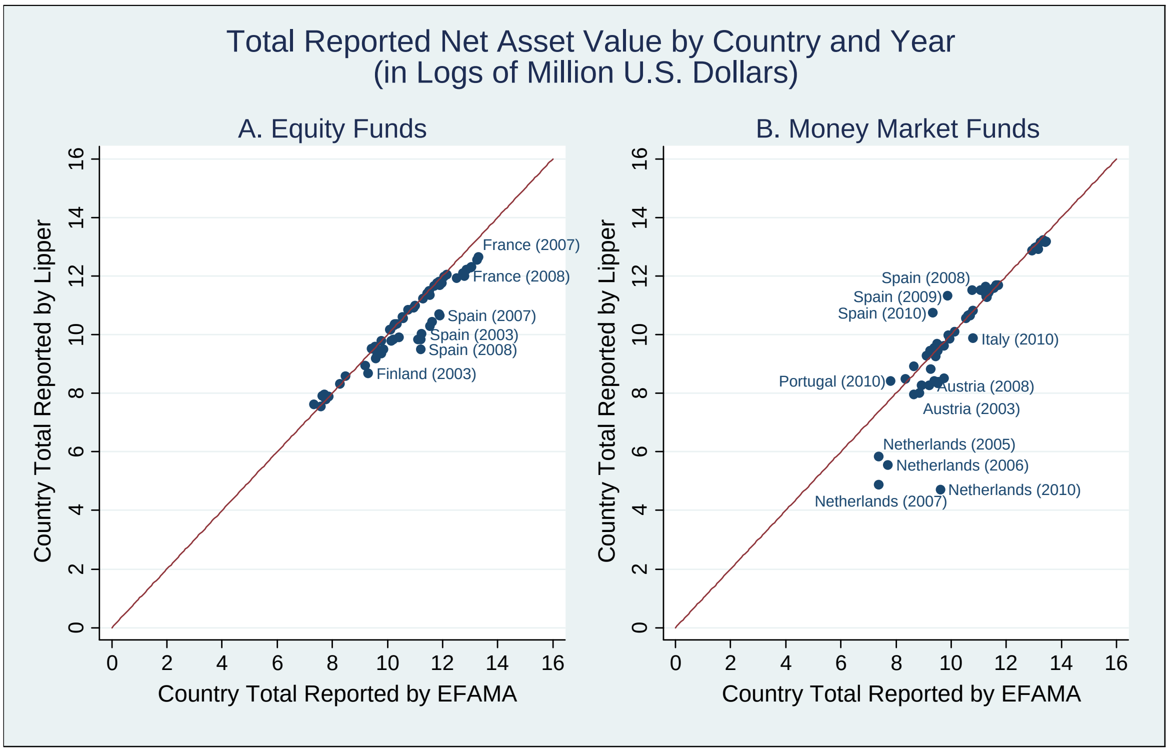

Figure 2: Plotted is the total net asset value (in the natural logarithm of million U.S. dollars) reported by the Lipper fund database on the y-axis against that reported by the European Fund and Asset Management Association (EFAMA) on the x-axis for the eight eurozone countries—Austria, Finland, France,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

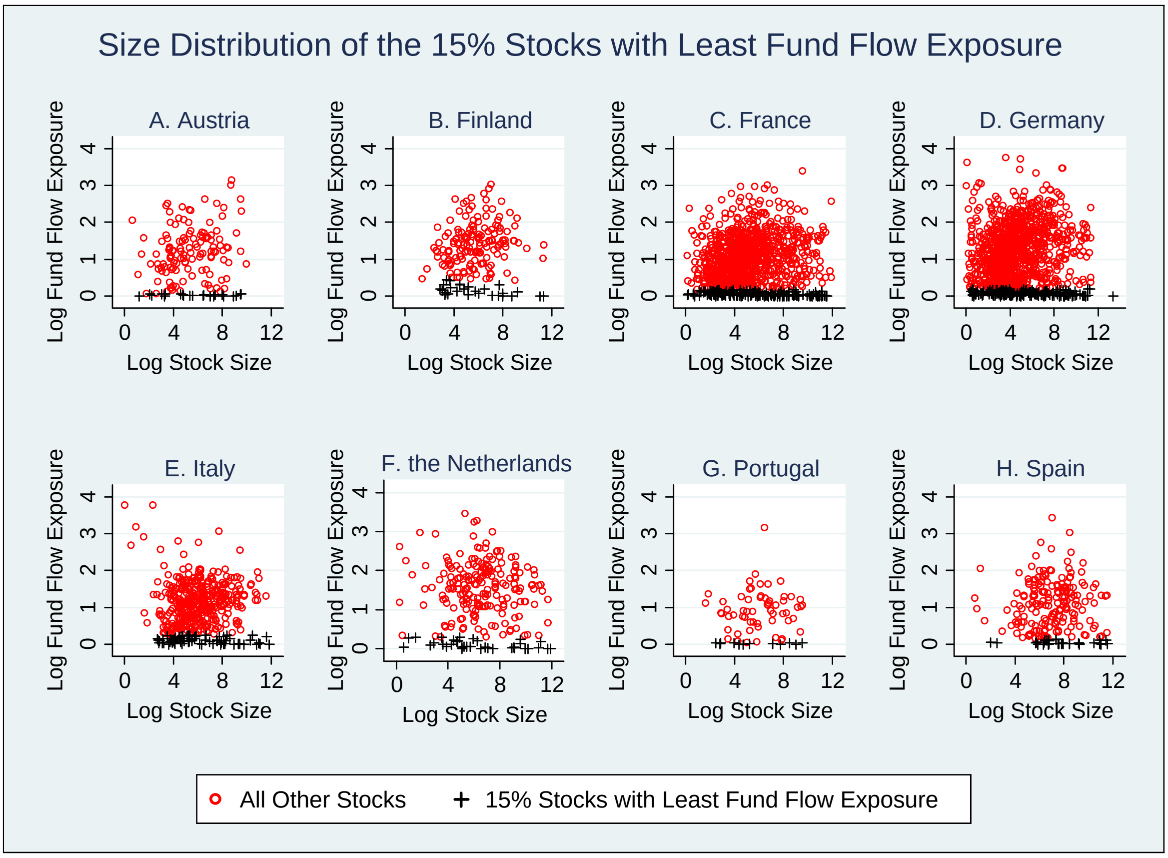

Figure 3: Plotted is the fund flow exposure for stocks in eight Eurozone countries against the stock size. The 15% of stocks with the lowest fund flow exposure in each country are marked by black crosses, whereas all other stocks are marked by red circles. Here we calculate the fund flow exposure for each stock as the natural logarithm of one plus the average (over the sample period 2003/1–2010/4) 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. The x-axis represents the natural logarithm of one plus the market capitalization value (in million U.S. dollars) of the stock, averaged over the same sample period.