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Daniela Andrén, Dan Gabriel Anghel and Olha Kyrkelan

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Örebro University School of Business
SE-701 82 Örebro, Sweden

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Daniela Andrén[1,2,3] Dan Gabriel Anghel[4,5] Olha Kyrkelan[4]

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[1]Örebro University School of Business, Örebro, Sweden
[2]Global Labor Organization (GLO)
[3]Central European Labour Studies Institute (CELSI)
[4]Faculty of Finance and Banking, The Bucharest University of Economic Studies, Bucharest, Romania
[5]Institute for Economic Forecasting, Romanian Academy, Bucharest, Romania

Abstract

Retail investors in emerging capital markets often hold narrow portfolios, but whether this reflects limited investment opportunities or investor-specific frictions remains an open empirical question. Using 96,220 investor-date observations for 43,360 clients of a large Romanian broker between December 2020 and August 2023, we examine portfolio breadth, concentration, and investment allocation. Individual portfolios are narrow: the median contains five distinct financial instruments, and the mean effective number of holdings is 4.25. Brokerage portfolio value is positively associated with portfolio breadth, both across and within investors, although the evidence does not establish a causal relationship. Conditioning on portfolio value reduces several demographic and socioeconomic differences, but substantial heterogeneity remains. Fund and ETF participation increases with portfolio value, while the unconditional portfolio share allocated to these instruments declines. Domestic concentration is substantial: approximately three-quarters of investors hold only Romanian-classified financial instruments in August 2023, and the within-investor domestic-share gradient is small. In contrast to narrow individual portfolios, the aggregate client portfolio has an effective number of approximately 44 holdings, while 3,534 distinct foreign financial instruments appear in client holdings over the observation period. These comparisons reveal a broad observed range of financial instruments but do not establish the opportunities effectively available to each investor. Finally, changes in investor composition substantially affect portfolio statistics: between December 2022 and August 2023, average portfolio breadth declines by 0.58 instruments, despite increasing by 1.19 among continuing investors. The findings provide new evidence on the relationships between portfolio value, investment allocation, and investor composition in an emerging capital market, without attributing these patterns to a uniquely identified mechanism.

Keywords: household finance; portfolio diversification; retail investors; emerging capital markets; Romania

JEL: G11, G51, G53, D14, P34

1 Introduction

Household participation in capital markets has expanded rapidly in the past decade, but the empirical literature that describes what households actually *hold* remains concentrated in a small number of high-income countries where administrative records and brokerage data have enabled detailed analyses of household portfolios — above all Sweden (Calvet et al., 2007, 2009; Betermier et al., 2017), and to a lesser extent Finland (Grinblatt and Keloharju, 2001; Grinblatt et al., 2011), the Netherlands (van Rooij et al., 2011; von Gaudecker, 2015) and the United States (Barber and Odean, 2000; Goetzmann and Kumar, 2008). Much of the empirical evidence on under-diversification, home bias, and the relationship between financial wealth and portfolio breadth therefore comes from deep, liquid markets with large asset menus, mature intermediation, and decades of accumulated participation. Whether these same regularities describe investors in a thin market with less developed capital-market institutions is an open empirical question. Romania's stock exchange lists fewer than 400 tradable equities, and a small number of firms account for most market capitalization. In such a setting, narrow portfolios may reflect limited domestic investment opportunities, investor-level constraints, or portfolio preferences. As retail participation expands, an additional question arises: whether changes in average portfolio characteristics reflect changes in existing investors' holdings or changes in the composition of the investor population.

This paper investigates portfolio breadth, concentration, and investment allocation among Romanian retail investors, examining both differences across investors and changes over time. We use proprietary, anonymized, position-level holdings for the entire active retail client base of a large Romanian broker at four observation dates: December 31, 2020; December 31, 2021; December 31, 2022, and August 22, 2023. The data cover 43,360 distinct investors and 766,637 positions, corresponding to 96,220 investor-date observations. Aggregate holdings rise from RON 2.15 billion in the first snapshot to RON 5.17 billion in the last. These observations allow us to examine portfolio breadth, concentration, fund participation, and domestic asset allocation, as well as their associations with portfolio value and investor characteristics. Throughout the paper, a financial instrument denotes a distinct traded instrument identified by its trading symbol in the brokerage records; a position denotes an investor's holding of that instrument at a given observation date. Portfolio breadth is the number of distinct financial instruments held by an investor, while the effective number of holdings measures the concentration of portfolio value across those instruments.

The data and the setting have three key features. First, holdings are observed at the position level rather than self-reported in a survey. Second, the panel spans the pandemic-era participation wave, during which the observed client base roughly quadrupled. The repeated snapshots allow us to distinguish changes among investors present at successive dates from changes resulting from the appearance and disappearance of investors in the observed sample. Third, the setting is a market that was re-established after the collapse

of the communist centrally planned economy and subsequently shaped by privatization, regulatory reforms, and the development of capital-market institutions.

Our contribution is threefold. First, we provide detailed evidence on retail portfolio breadth, concentration, fund and ETF ownership, and domestic allocation in an emerging capital market, documenting how these characteristics vary with brokerage portfolio value and investor demographics. Second, we use repeated holdings observations to examine changes within investors and compare individual portfolios with the substantially broader range of financial instruments observed across the brokerage client base. Third, we decompose changes in average portfolio characteristics into changes among continuing investors and changes in the composition of the observed investor population. This decomposition reveals that average portfolio breadth and the effective number of holdings can decline in the full cross-section even as both increase among continuing investors. Together, these analyses provide evidence on how retail portfolio characteristics differ across investors and evolve as participation expands, without attributing the observed patterns to a uniquely identified economic mechanism.

Household-finance theories of information costs, familiarity, and delegated investment guide the analysis, but the empirical design does not identify these mechanisms causally. We begin with the frictionless mean–variance benchmark and introduce three frictions that are standard in the household-finance literature. The first is information or attention costs associated with acquiring and maintaining distinct financial instruments, which can imply a finite optimal portfolio breadth that increases with financial wealth under appropriate assumptions (Merton, 1987; Van Nieuwerburgh and Veldkamp, 2010). The second is delegation through investment funds and exchange-traded funds (ETFs), which allows investors to obtain diversification through pooled investment vehicles without holding many instruments directly (Campbell, 2006). The third is informational asymmetry or familiarity, which may tilt portfolios toward domestic and locally known assets (Coval and Moskowitz, 1999; Huberman, 2001; Van Nieuwerburgh and Veldkamp, 2009). These mechanisms may operate jointly and need not generate mutually exclusive empirical predictions. They motivate the relationships examined in the empirical analysis rather than providing mutually exclusive explanations that the data can distinguish.

The framework yields four hypotheses. The first three concern observable relationships between investor characteristics, portfolio value, and portfolio composition; the fourth evaluates the range of financial instruments held across the broker's client base relative to individual portfolio concentration. The corresponding empirical assessments differ in their statistical interpretation and ability to distinguish competing explanations. We do not estimate structural parameters and we do not claim causal identification.

Evidence from the platform indicates that the range of instruments held across the client base is substantially broader than the portfolios held by individual investors. This finding is difficult to reconcile with a simple explanation based exclusively on the number of

instruments observed in the broker's client holdings. In August 2023, clients collectively held 3,160 distinct financial instruments, and the value-weighted portfolio formed by aggregating all client holdings had an effective number of holdings of approximately 44, compared with a mean effective number of 4.11 across individual portfolios. However, this comparison does not establish the set of instruments effectively accessible to each investor or the costs associated with holding them.

Foreign financial instruments are also held by some investors in the broker's client base, demonstrating that access to foreign instruments is not universally absent. At least 3,534 foreign instruments were held by someone in the client base across the four observation dates, including 1,838 foreign equities and 335 foreign exchange-traded funds. Yet the median investor holds no foreign financial instrument at any of the four observation dates, and 74% of investors hold only Romanian-classified financial instruments in August 2023. These findings illustrate the coexistence of extensive foreign-instrument holdings across the client base and predominantly domestic portfolios among individual investors. They are compatible with several explanations, including familiarity, information costs, differences in effective investment access, and investor preferences, but do not distinguish their relative importance. The first set of results concerns the relationship between portfolio value and portfolio breadth. The number of distinct instruments increases with portfolio value across investors. Conditioning on portfolio value substantially reduces several differences associated with age, education, and income, although sizable conditional differences remain for some categories. Women also hold approximately 1.4 fewer distinct financial instruments than men among investors with similar brokerage portfolio values. The positive association also appears in investor fixed-effects estimates. However, a separate instrumental-variable analysis using passive portfolio returns does not establish a statistically significant causal response of breadth to changes in portfolio value. These results document a robust association while leaving its causal interpretation unresolved. Second, investors rely on funds and exchange-traded funds in a way that depends on portfolio size. The unconditional share of portfolio value held in these instruments falls with portfolio size, while the probability of holding at least one fund or exchange-traded fund rises. These findings describe distinct relationships at the allocation and participation margins of fund ownership and can be evaluated against theoretical predictions concerning delegated investment. However, the combined category includes investment vehicles with different underlying exposures and investment objectives, so observed ownership cannot be interpreted directly as the acquisition of diversified underlying assets. Third, domestic portfolio allocation remains substantial across investors and observation dates. The median investor holds no foreign-classified financial instrument, and the standardized within-investor relationship between brokerage portfolio value and the domestic portfolio share is considerably smaller than the corresponding relationship for portfolio breadth. This indicates that domestic allocation changes relatively little as portfolio value varies within investors, although the data do not establish

why this pattern arises.

The apparent deterioration of average portfolio breadth in 2023 reflects changes in the sample despite increases in average breadth among continuing investors. Mean portfolio breadth falls between the 2022 and 2023 observation dates because investors observed in August 2023 but absent from the preceding snapshot hold substantially narrower portfolios than continuing investors. Continuing investors add increase their mean portfolio breadth by 1.18 holdings on average, but 15,230 investors appearing in the later snapshot hold only about 4.5 holdings on average, which contributes to the decline in the cross-sectional mean. Specifically, mean portfolio breadth falls by approximately 0.58 financial instruments between December 2022 and August 2023, while increasing by approximately 1.19 among investors present at both dates.

The same composition effect explains most of the increase in the equally weighted mean investor-level domestic portfolio share between December 2022 and August 2023. Entrants have a mean domestic portfolio share of approximately 96%, compared with 91% for continuing investors. As a result, cross-sectional measures of retail portfolio breadth, concentration, and domestic allocation in a fast-growing client base can be substantially affected by changes in the composition of the observed investor population.

The paper relates to four strands of research. First, it adds to the household-finance tradition surveyed by Campbell (2006), Guiso and Sodini (2013), and Gomes et al. (2021) by examining portfolio structure and its evolution within a rapidly expanding retail investor population in an emerging capital market. Administrative holdings data are increasingly available outside high-income countries. For example, Campbell et al. (2014) study a large representative sample of Indian retail equity investors drawn from depository records. Our analysis contributes to this literature by providing detailed brokerage holdings evidence from Romania, a post-transition Central and Eastern European market that has received comparatively limited attention in investor-level household-finance research.

Second, it contributes to the literature on under-diversification (Blume and Friend, 1975; Goetzmann and Kumar, 2008; Calvet et al., 2007; Polkovnichenko, 2005; von Gaudecker, 2015) by examining the portfolio-value gradient in the number of holdings in a setting where the domestic investment opportunity set is much smaller than in the markets where this gradient was first established. We do not compare the magnitude of our estimates with earlier studies because the objects measured are different. Calvet et al. (2007), for example, evaluate investment efficiency for the full household portfolio and jointly with the participation decision. In contrast, we count and weight holdings within a single brokerage account in a clientele that is predominantly invested in Romanian-classified financial instruments. What is comparable is the sign and general pattern of the portfolio-value gradient and its role in organizing the cross-section. The analysis also distinguishes changes in portfolio breadth among continuing investors from changes in the average breadth of the observed investor population, extending the descriptive evidence beyond static portfolio comparisons.

Third, it speaks to the home-bias literature (French and Poterba, 1991; Coval and Moskowitz, 1999; Van Nieuwerburgh and Veldkamp, 2009). The domestic portfolio-value share is high, with the median investor holding only Romanian-classified financial instruments, exhibiting relatively limited variation across portfolios containing up to 20 distinct financial instruments and a small standardized within-investor association with portfolio value. This pattern is consistent with familiarity and fixed information costs, although the available evidence does not distinguish these explanations from other determinants of domestic asset allocation.

Fourth, it relates to work on capital markets in transition economies (Claessens et al., 2000; Pistor et al., 2000; Estrin et al., 2009; Pop et al., 2015; Pop, 2020) by documenting contemporary retail portfolio choices in an institutional environment shaped by central planning, post-communist privatization, and the subsequent development of capital-market institutions. The analysis complements research on the institutional evolution of transition economies by providing investor-level evidence on portfolio holdings within this setting, without attributing present-day investment choices causally to specific historical reforms.

We are explicit about what these data cannot do. The sample covers clients of a single broker. We treat it as a census of that client base rather than as a random sample of Romanian households, and discuss selection in Section 4. The observations are end-of-day snapshots rather than transactions, so the data describe holdings, not trades. Background characteristics are recorded once at account opening, and their availability varies systematically with observed portfolio characteristics and observation dates. We treat patterns of missing information as an object of study in Section 5.6. Finally, our measures count and weight holdings, so they describe the number and concentration of holdings, not portfolio risk. For example, two global exchange-traded funds can provide broader diversification than ten domestic bank stocks. We state this limitation explicitly in Section 3.1 before presenting results. These measurement and sampling constraints also limit our ability to identify the mechanisms underlying observed portfolio choices or to generalize the findings to the Romanian household population.

The remainder of the paper is organized as follows. Section 2 describes the institutional development of Romania's capital market and the investment environment relevant to retail investors. Section 3 sets out the conceptual framework and hypotheses. Section 4 describes the data, measures, and sample conventions. Section 5 reports the results, including the decomposition of changes in cross-sectional portfolio characteristics, and Section 6 interprets them and discusses policy implications and limitations. Section 7 concludes.

2 Institutional setting

The domestic listed market, the set of financial instruments accessible through the broker, and the investment opportunities effectively available to each individual investor are three different concepts. The institutional evidence describes the first two. The holdings data reveal which financial instruments investors actually hold. Neither source, by itself, establishes the complete feasible investment set of every investor.

Portfolio choice depends on the financial instruments available to an investor and the conditions under which they can be held. Before describing what Romanian retail investors hold, we describe the domestic investment opportunity set and access to foreign financial instruments through the observed brokerage platform, and why this opportunity set has its current shape.

Five institutional features jointly shape the domestic investment opportunity set. These are: a market history marked by closure under communism and re-establishment during the transition to a market economy; pyramid and fraudulent investment schemes during the early transition; a privatization legacy of dispersed but largely passive ownership; high sectoral and ownership concentration with limited free float; and the recent expansion of exchange-traded government bonds offered directly to retail investors. A sixth feature, low-cost platform access to foreign markets, expands the range of financial instruments accessible to the broker's clients beyond those listed domestically, although it does not eliminate investor-specific costs and constraints.

2.1 Closure, re-establishment and the privatization legacy

Organized securities trading in Romania began on 1 December 1882, following a royal decree under King Carol I. The institution operated as a combined stock and commodities exchange until it was closed by the communist regime in 1948 (Bucharest Stock Exchange, 2022). The closure dismantled not only the exchange but also much of the supporting infrastructure of a securities market, including intermediation, disclosure practices, and the legal and administrative routines needed for trading and investor protection. Under state socialism, capital allocation was administrative, and market-based price discovery, external equity finance, and tradable ownership largely disappeared.

This break matters for the modern market. When the exchange was re-established, it had to be built largely from scratch in an economy where private ownership, securities regulation, and corporate governance were still being created (Pistor et al., 2000; Berkowitz et al., 2003). Law No. 52/1994 established the legal basis (Parliament of Romania, 1994). The Bucharest Stock Exchange reopened on 20 November 1995 with six listed companies and 45 trades (Bucharest Stock Exchange, 2020a). The long absence of household experience with securities may therefore shape present-day participation and trust, consistent with

evidence that macroeconomic and political experiences affect risk-taking (Malmendier and Nagel, 2011; Alesina and Fuchs-Schundeln, 2007) and that trust and social capital are key determinants of stock-market participation (Guiso et al., 2004, 2008).

Credibility was further undermined by pyramid schemes during the early transition period. The most prominent was *Caritas*, which attracted millions of participants before collapsing in 1993–1994 (Verdery, 1995; Biberaj et al., 1997). Similar schemes also appeared across Eastern Europe under weak regulation and limited financial experience (Gogozan, 2009; Harison and Mihály, 2021).

These episodes imposed large household losses and likely reduced trust in financial institutions and market-based investing. Lower trust can discourage stock-market participation by reducing households' willingness to hold risky financial assets (Biberaj et al., 1997; Gurun et al., 2018).

The early structure of the Romanian market was shaped less by new share issues than by mass privatization. The Mass Privatization Program of 1995–1996 distributed vouchers to adult citizens, who could exchange them for shares in state-owned enterprises (OECD, 2022). This voucher approach has been analysed for the region by Boycko et al. (1994), Megginson and Netter (2001), and Estrin et al. (2009).

The program created many shareholders but not necessarily many active investors. It brought to market firms that often had limited incentives to meet the disclosure and governance standards associated with active public trading. The OECD reports that roughly nine million accounts in the Central Securities Depository associated with privatization-era allocations were still inactive as of end-2018, with an estimated value of about EUR 2.2 billion (OECD, 2022). Ownership that is widespread but largely inactive can contribute to limited trading and governance challenges. Consequently, the number of listed firms is not a good guide to how many securities are actively tradable.

The RASDAQ market, launched in November 1996, was created to absorb the roughly 5,000 companies that privatization had turned into issuers (OECD, 2022). A large number of listings did not translate into a deep market with active trading and strong disclosure.

By the end of 2003, 62 companies were listed on the Bucharest Stock Exchange (BVB) and 2,460 traded on RASDAQ. Combined market capitalization reached 10.7% of GDP, but the aggregate turnover ratio was only 7%. Trading and market value were highly concentrated. The ten largest BVB firms accounted for 87% of BVB capitalization, and more than 81% of trading volume was concentrated in five Financial Investment Companies (SIFs), two banks, and Petrom (Shojai and Noel, 2005).

Pop et al. (2015) argue that RASDAQ mainly facilitated ownership consolidation by existing shareholders rather than serving as a venue for portfolio investment. RASDAQ was wound down by Law No. 151/2014 (Parliament of Romania, 2014; Tuca Zbarcea and Asociatii, 2014). Firms were required to migrate to the Main Market, to the newly created AeRO segment for small and medium-sized enterprises (SMEs) (Bucharest Stock Exchange,

2014, 2015), or to delist.

The Sibiu derivatives exchange was absorbed in 2016–2017 (Bucharest Stock Exchange, 2016), urther consolidating organized securities trading under the BVB.[1]

For household portfolios, the key point is that today's concentration is not a new development. Since the market's re-establishment, a formally broad listed universe has coexisted with a narrow, actively traded core.

2.2 Size, concentration and free float

Figure 1 places Romania in regional perspective relative to the Central and Eastern European countries used by Pop (2020). The comparisons extend through 2024 and provide institutional context beyond the 2020–2023 period covered by the investor-level data. Three features are relevant. First, market scale differs sharply across countries. In 2024, Poland's domestic market capitalization is US$197 billion with 750 listed domestic firms. Romania has US$47 billion and 83 listed domestic firms. Both panels use a logarithmic vertical scale because a linear scale compresses the smaller markets and hides meaningful differences. Second, Romania has converged toward Hungary in market value. Romanian capitalization, below Hungary's for most of the 2000s and 2010s, overtook it in 2021 and has remained above it since. The large increase in 2023, a 65% rise in one year, coincides with Hidroelectrica's listing in July 2023. At EUR 1.9 billion, it was the largest initial public offering in the Bucharest Stock Exchange's history (Bucharest Stock Exchange, 2023). That a single issuer is visible in the country-level series is consistent with the substantial influence that a major listing can have on a relatively small equity market. Third, Romania's market has deepened in value without broadening in the number of issuers. Poland's listed universe expanded from 211 firms in 2004 to a peak of 861 in 2016 and then consolidated to 750, aided by its NewConnect market for small and medium-sized enterprises. Romania's count rose from 55 firms in 2004 and has been essentially flat at 80–86 firms since 2013.

However, Romania's international classification remains mixed. FTSE Russell promoted Romania from Frontier to Secondary Emerging status effective September 2020 (FTSE Russell, 2020; Bucharest Stock Exchange, 2020b). MSCI, by contrast, continues to classify Romania as a Frontier market (MSCI Inc., 2025). This split classification may affect Romania's inclusion in different international equity indices and the associated flows from index-tracking investors.

Market capitalization does not directly measure the range of financial instruments that

[1] SIBEX (Sibiu Stock Exchange) was established in 1994. It initially operated as a commodities exchange and later developed mainly as Romania's derivatives exchange, offering futures and options on equities, stock indices, currencies, and commodities. It also operated a regulated spot market and an alternative trading system for equities, but both remained small relative to the BVB. Derivatives trading peaked in 2006 at over 4.2 million futures and options contracts, then declined. At the time of the merger, 24 SIBEX companies were transferred to the BVB's AeRO Base segment and 12 to AeRO Standard; none transferred to the BVB's main market (Bucharest Stock Exchange, 2018).

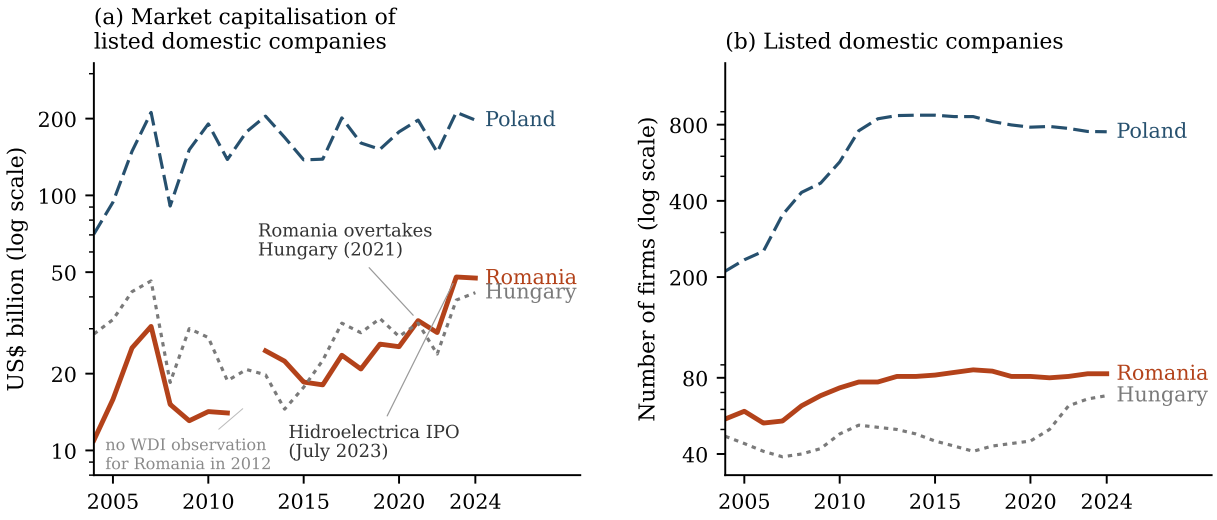


Figure 1 Romania in regional perspective, 2004–2024

Notes: Authors' own figure. Underlying data: World Bank, World Development Indicators, series `CM.MKT.LCAP.CD` (market capitalization of listed domestic companies, current US$) and `CM.MKT.LDOM.NO` (listed domestic companies, total), retrieved 10 August 2026. Both panels use a logarithmic vertical scale. The Romanian capitalization series has no WDI observation for 2012; the line is broken rather than interpolated. Listed-firm counts are unavailable before 2004.

investors can feasibly use for portfolio diversification. In several of the largest issuers, the Romanian state and other strategic holders retain majority stakes, which limits the free float. The OECD documents that within the BVB Main Market the average free float of stocks that had not traded in the previous week was almost half that of stocks that had traded (OECD, 2022, 2023).

Sectoral concentration reinforces this. In July 2024, the BVB reported that energy, utilities, and financial companies accounted for about 58% of regulated-market capitalization excluding investment funds (Bucharest Stock Exchange, 2024). Limited free float and concentration in a small number of large issuers imply that the number of independent domestic exposures cannot be inferred from the number of listings alone. This configuration is central to classic work on illiquidity and asset pricing (Amihud and Mendelson, 1986; Amihud, 2002) and is common in emerging markets (Bekaert and Harvey, 2003; Karolyi, 2015).

Bank lending continues to dominate corporate finance. The OECD reports that about 99% of Romanian companies' debt financing comes from banks, and over the period beginning in 2008 covered by its report, the Main Market has recorded 45 delistings against 16 new listings (OECD, 2022). The domestic equity universe has therefore been shrinking at the extensive margin even as market capitalization has grown.

A distinctive recent change in the opportunity set is the state's expansion of exchange-traded government bond issuance for retail investors. Government bonds issued by the Ministry of Finance and listed on the BVB under the retail-focused *Fidelis* program are low-denomination, tax-advantaged, exchange-traded financial instruments designed for households. In our data this appears as a supply expansion. The number of distinct government-

bond lines held by clients rises from 6 in 2020 to 31 in 2023, and by 2023 one such line (R2408A) is the largest financial instrument by value in the aggregate client portfolio. This policy-driven change matters for portfolio choice because it offers households a nominally safe and exchange-traded financial instrument at a time when the real return on deposits turned sharply negative.

2.3 Access to foreign assets and the participation wave

The broker whose clients we observe is a large domestic full-service intermediary. It routes orders directly to market and does not internalise or trade against clients. It is subject to the Markets in Financial Instruments Directive (MiFID) framework and the national investor-compensation scheme, holds client funds in segregated accounts, and offers no contracts for difference (CFDs) or comparable leveraged retail products. Importantly, it provides access to both the domestic exchange and major international markets through a single platform, with local tax withholding and reporting and Romanian-language support.

This matters for interpretation. In this setting, limited platform access is unlikely to be the main barrier separating Romanian households from foreign assets. However, access through the platform does not imply that investors face identical information costs, transaction costs, or practical investment opportunities.[2]

Consistent with this, at least 3,534 distinct foreign financial instruments were held by clients of this broker during our sample, including 1,838 foreign equities and 335 foreign exchange-traded funds. Foreign financial instruments are therefore accessible through the platform and held by some investors, but not by the median investor in any year.

The period we observe is also not a steady state. The client base of this broker roughly quadrupled between end-2020 and August 2023, from 9,837 to 41,234 active accounts, in line with the pandemic-era retail wave documented in other markets (Barber et al., 2022; Welch, 2022). Cross-sectional statistics in such a setting combine investor cohorts with very different portfolio sizes and tenures. Section 5.6 makes this explicit and shows that it is decisive for interpreting time trends.

[2]International investing by Romanian residents was liberalised in stages. A key change took effect on 1 January 2003, when National Bank of Romania Circular No. 26/2001 removed the prior-authorization requirement for residents' investments in foreign securities, as part of capital-account liberalisation ahead of European Union accession (National Bank of Romania, 2001; European Commission, 2003; International Monetary Fund, 2003). The framework for cross-border investment services became fully operational upon European Union accession on 1 January 2007 under Law No. 297/2004. By the 2020s, access to international financial markets through both domestic and foreign intermediaries was well established.

2.4 Market microstructure, liquidity, volatility and informational efficiency

Trading on the Bucharest Stock Exchange is electronic, but liquidity is concentrated in a relatively small set of actively traded stocks (Anghel et al., 2020). Mass privatization created many shareholders, and many did not trade regularly, contributing to thin trading in large parts of the market.

Low liquidity, high transaction costs, and wide bid–ask spreads have therefore been important frictions (Dragotă and Mitrică, 2004; Anghel, 2017). Short selling is permitted in principle, but in practice it remains largely unavailable because there is no active securities-lending market and liquidity is limited. Information asymmetries also appear persistent, with informed trading linked to macroeconomic conditions, liquidity, and volatility (Cepoi and Toma, 2016).

Studies of market efficiency point to meaningful time variation. Early evidence documents predictable price dynamics and departures from fundamental valuation (Dragotă and Mitrică, 2004). Cross-country studies also find deviations from weak-form efficiency in Romania, similar to other post-communist markets (Dragotă and Ţilică, 2014). Anghel (2015) finds that efficiency varies over time and does not improve monotonically, and Anghel (2018) suggests that investor sentiment can temporarily affect returns.

At higher frequencies, evidence points to autocorrelation, intraday seasonality, and other departures from a random-walk benchmark (Anghel, 2017). However, statistical predictability does not generally translate into systematic abnormal profits once transaction costs, trading restrictions, and data-snooping concerns are taken into account. The cited studies report retail trading costs of approximately 0.2% to 1% per trade or the periods and trading arrangements they examine (Anghel, 2015, 2017; Anghel et al., 2020).

Overall, the Romanian market combines limited and uneven liquidity with substantial information frictions. These features affect both the attractiveness and the practical tradability of individual securities, and they may therefore shape retail portfolio choices. They also suggest why the number of listed financial instruments may not fully capture the investment opportunities that individual investors can effectively exploit.

3 Theoretical framework and hypotheses

3.1 Measurement

Let w_{ijt} denote the value share of financial instrument j in investor i's portfolio at observation date t, with $\sum_{j=1}^{N_{it}} w_{ijt} = 1$, and let V_{it} denote the value of that portfolio in Romanian leu. Portfolio breadth is N_{it}, the number of distinct financial instruments held. Concentration is measured by the Herfindahl–Hirschman index of portfolio weights, $HHI_{it} = \sum_j w_{ijt}^2$,

and by its reciprocal, the effective number of holdings,

$$N_{\text{eff},it} \;=\; \left(\sum\nolimits_j w_{ijt}^2\right)^{-1}, \tag{1}$$

the number of equally weighted financial instruments that would deliver the observed concentration.

For completeness we also report normalized Shannon entropy, $H_{it}^* = -\sum_j w_{ijt} \log w_{ijt} / \log N_{it}$, and the Gini coefficient of portfolio weights in Appendix D, but we do not use them in the main text. Because entropy divides by $\log N_{it}$, it measures balance conditional on breadth rather than diversification and is undefined for single-asset portfolios. The Gini coefficient of weights equals zero both for a single-asset portfolio and for an equally weighted one. Portfolio value V_{it} and portfolio weights w_{ijt} are distinct objects throughout.

These statistics measure breadth and concentration across financial instruments and ignore the covariance structure. Three consequences follow. First, may understate the diversification of underlying exposures in portfolios containing broadly diversified investment funds or ETFs: two positions in global exchange-traded funds count as $N = 2$ while potentially providing exposure to thousands of underlying securities. Without look-through data, every fund position is counted as a single financial instrument. The resulting measures cannot establish the extent of economic risk diversification obtained through funds. Second, they may overstate the economic diversification provided by concentrated domestic equity: ten Bucharest-listed issuers, five of them state-linked energy and utility firms exposed to a common regulatory factor, may provide less diversification of portfolio risk than ten randomly drawn global equities. Third, they describe holdings at one broker rather than a household balance sheet, so an investor holding a single Romanian equity here may be well diversified overall, and the converse also holds (Guiso et al., 2002; Badarinza et al., 2016). We retain these measures because they are standard in the literature to which our results speak (Blume and Friend, 1975; Calvet et al., 2007; Goetzmann and Kumar, 2008), while interpreting them strictly as measures of portfolio breadth and concentration within the observed brokerage holdings.

3.2 Theory and hypotheses

In the frictionless mean–variance benchmark (Markowitz, 1952), portfolio choice depends on expected returns, return covariances, and investor preferences. Under additional equilibrium assumptions, investors may hold a common portfolio of risky assets, with differences in overall risk exposure reflecting differences in risk tolerance. Internationally, a fully diversified global market portfolio would assign domestic securities a weight corresponding to their share of global market capitalization, a small fraction of one percent for Romania (French and Poterba, 1991). We do not test this benchmark. Observed household

portfolios frequently depart from its restrictive predictions, and its role here is to organize candidate explanations of observed narrowness into (i) the size of the feasible menu, (ii) the cost of an additional position, (iii) heterogeneous information, and (iv) non-mean–variance preferences. The first is a supply-side explanation specific to a thin emerging market; the remainder are demand-side. The hypotheses below are chosen so that the data can be used to examine selected observable implications of these explanations. Each is stated so that the theoretical claim is the alternative and the null is the absence of the effect it predicts.

A fixed cost per financial instrument is the departure of most relevance. In Merton (1987), an investor holds only the assets she recognizes, and recognition is costly. In Van Nieuwerburgh and Veldkamp (2010), investors acquire information about a limited set of assets and then over-weight them, so that under-diversification is the outcome of rational information choice.

These models motivate the possibility that investors acquire information about, or devote attention to, only a subset of available financial instruments. When expanding portfolio breadth involves fixed costs, the benefits of holding additional instruments may be more substantial for investors with larger portfolios. The relationship between portfolio value and the optimal number of holdings nevertheless depends on the specification of preferences, information costs, and investment opportunities. The approximately log-linear relationship examined below is therefore an empirical specification motivated by this literature, rather than a direct implication established by the cited models.

The same characteristics that predict portfolio value also predict the number of holdings. This raises an additional empirical question: how much of the observed association between investor characteristics and portfolio breadth remains after conditioning on portfolio value? The answer provides evidence on cross-sectional heterogeneity, but does not establish that portfolio value mediates the effects of those characteristics.

Hypothesis 1 (portfolio breadth and portfolio value). H1a: Portfolio breadth, measured as the number of distinct financial instruments held, is positively associated with portfolio value. An approximately log-linear relationship is examined empirically. H1b: Differences in portfolio breadth across demographic and socioeconomic groups are assessed both before and after conditioning on portfolio value. The extent to which these differences diminish is interpreted descriptively rather than as evidence of causal mediation.

H1a is tested with

$$N_{it} \;=\; \alpha \;+\; \beta \ln V_{it} \;+\; \delta_t \;+\; \varepsilon_{it}, \tag{2}$$

where δ_t are date fixed effects and standard errors are clustered by investor. The null is $H_0 : \beta = 0$ against the one-sided alternative $H_1 : \beta > 0$. Equation (2) is also estimated with $\log N_{\text{eff},it}$ as the dependent variable. We also inspect decile profiles of N_{it} and $N_{\text{eff},it}$ for departures from monotonicity and log-linearity. A corollary is that if investors add new instruments at small weights, N_{eff} grows more slowly than N, so the ratio $N_{\text{eff},it}/N_{it}$ declines

with V_{it}. If additional financial instruments are held at relatively small portfolio weights, portfolio breadth may increase more rapidly than the effective number of holdings. We therefore also examine how $N_{\text{eff},it}/N_{it}$ varies with portfolio value.

H1b is assessed by comparing

$$N_{it} \;=\; \alpha \;+\; \textstyle\sum_k \gamma_k D_i^k \;+\; \delta_t \;+\; u_{it} \tag{3}$$

with

$$N_{it} \;=\; \alpha \;+\; \textstyle\sum_k \tilde{\gamma}_k D_i^k \;+\; \beta \ln V_{it} \;+\; \delta_t \;+\; u_{it}, \tag{4}$$

where D_i^k are indicators for the categories of one account-opening characteristic. If demographic differences in the number of holdings mainly reflect differences in portfolio size, then adding $\ln V_{it}$ should shrink the coefficients on D_i^k. We therefore test $H_0 : \tilde{\gamma}_k = \gamma_k$ against $H_1 : |\tilde{\gamma}_k| < |\gamma_k|$. The quantity $1 - \tilde{\gamma}_k/\gamma_k$ reports the share of the raw gap absorbed. We compare the coefficients across specifications to assess how the conditional demographic differences change when portfolio value is included. Where the unadjusted coefficients are sufficiently different from zero, $1 - \tilde{\gamma}_k/\gamma_k$ summarizes proportional attenuation. This statistic is descriptive and is not interpreted as the proportion of a causal effect mediated by portfolio value.

Equation (2) uses variation across investors, so it may partly reflect sorting. Investors who prefer broader portfolios may also be those who accumulate larger portfolios. A within-investor specification is therefore also reported,

$$N_{it} \;=\; \alpha_i \;+\; \beta^W \ln V_{it} \;+\; \delta_t \;+\; \nu_{it}, \tag{5}$$

which removes all time-invariant investor characteristics. This provides a stricter comparison, but it is not an identification strategy because $\ln V_{it}$ remains jointly determined with portfolio composition within investors. In particular, net investments, withdrawals, portfolio rebalancing, and price changes may jointly affect observed portfolio value and breadth.

Investment funds and exchange-traded funds allow investors to obtain exposure to portfolios of underlying securities in a single holding. They therefore may reduce the need to acquire and monitor numerous financial instruments directly. Delegation should be particularly attractive for small portfolios, where the fixed per-holding cost c is large relative to invested wealth (Campbell, 2006; Gomes et al., 2021). von Gaudecker (2015) documents important relationships between financial literacy, advice, and household portfolio diversification, although these relationships do not uniquely identify fund ownership as the underlying mechanism. Participation may still rise with wealth if learning about these instruments financial instruments is itself costly.

Hypothesis 2 (fund participation and allocation). The relationship between portfolio value and delegated investment is examined along two dimensions: participation in investment

funds or ETFs and the share of portfolio value allocated to these financial instruments. We test whether fund participation increases with portfolio value while the portfolio share allocated to funds decreases. This joint pattern is an empirical implication to be evaluated rather than a unique prediction of delegated-diversification models.

Write f_{it} for the share of investor i's portfolio value held in investment funds and exchange-traded funds. We estimate two margins:

$$f_{it} = \alpha + \beta_f \ln V_{it} + \delta_t + \varepsilon_{it}, \qquad \mathbf{1}\{f_{it} > 0\} = \alpha + \beta_p \ln V_{it} + \delta_t + \eta_{it}, \quad (6)$$

where the second equation is a linear probability model for participation. Because H2 concerns both the portfolio-allocation margin ($\beta_f < 0$) and the participation margin ($\beta_p > 0$), we test it as a joint restriction using the joint cluster-robust covariance of $(\hat{\beta}_f, \hat{\beta}_p)$.

We report a Wald test of the point null $H_0 : \beta_f = \beta_p = 0$, but the relevant prediction is directional: $\beta_f < 0$ and $\beta_p > 0$. We therefore also use an intersection–union test (Berger, 1982), which rejects only if both one-sided tests reject at the 5% level in the predicted directions. This requirement establishes whether the two directional associations are jointly supported by the data, without identifying the mechanism generating them.

Home bias is a robust empirical regularity (French and Poterba, 1991). No single explanation accounts for its full extent (Ardalan, 2019), but many proposed mechanisms are informational. Investors may feel more informed about local assets (Coval and Moskowitz, 1999; Ivkovic and Weisbenner, 2005), invest in what is familiar (Huberman, 2001), and specialize in assets they already follow (Van Nieuwerburgh and Veldkamp, 2009). Distance, language, and culture also predict holdings (Grinblatt and Keloharju, 2001). Hedging motives exist but are typically second-order (Massa and Simonov, 2006). In an emerging market, these mechanisms may be reinforced by currency conversion, withholding taxes, reporting requirements, and low trust (Guiso et al., 2008; Giannetti and Koskinen, 2010).

A competing explanation is based on access or scale. If foreign investing is hard or not worthwhile for small portfolios, then the domestic share may fall as an investor's portfolio grows. However, information costs, familiarity, and access-related constraints may operate simultaneously and may generate similar relationships between portfolio value and domestic asset allocation.

Testing for "no response" is difficult because a small coefficient can reflect limited power rather than a true zero. We therefore compare the within-investor response of the domestic share with the within-investor response of the number of holdings. Let σ_N and σ_s denote within-investor standard deviations of N_{it} and s_{it}, and define standardized responses $\eta_N = \beta^W/\sigma_N$ and $\eta_s = \beta^W_s/\sigma_s$. This comparison indicates whether the two portfolio characteristics exhibit different degrees of association with within-investor changes in portfolio value. However, neither access-related constraints nor informational familiarity implies a particular ratio between these standardized coefficients without additional assumptions. We

therefore use this comparison to characterize the empirical patterns rather than to discriminate conclusively between mechanisms.

Hypothesis 3 (domestic portfolio allocation). The domestic portfolio-value share exhibits a smaller within-investor association with portfolio value than portfolio breadth, when both responses are standardized by their within-investor variability : $|\eta_s| < \eta_N$. This relationship is assessed empirically and interpreted as potentially consistent with familiarity and information-related frictions, without treating it as a unique test of those mechanisms.

Let s_{it} denote the domestic value share. The cross-sectional and within-investor gradients are

$$s_{it} = \alpha + \beta_s \ln V_{it} + \delta_t + \varepsilon_{it}, \qquad s_{it} = \alpha_i + \beta_s^W \ln V_{it} + \delta_t + \nu_{it}. \tag{7}$$

The two equations are estimated jointly on the same investors. We test $H_0 : \eta_N - |\eta_s| = 0$ against $H_1 : \eta_N - |\eta_s| > 0$ using the delta method and the joint cluster-robust covariance of $(\hat{\beta}^W, \hat{\beta}_s^W)$. As a robustness check, we also assess whether β_s^W is statistically equivalent to zero using two one-sided tests (Schuirmann, 1987) against a margin of one percentage point of domestic share per log point of portfolio value. The median domestic share by observation date and the cross-sectional estimate β_s are reported as descriptive support. The standardized-coefficient comparison and equivalence test address different empirical questions; neither separately identifies the mechanism underlying domestic portfolio concentration.

The remaining question is whether narrow portfolios reflect a binding opportunity set. Cross-country comparisons confound the menu with other cross-country differences. Observing an entire client base permits an internal assessment, posed as a comparison between the financial instruments held across the platform and those held by individual investors. Aggregating all clients' holdings into one value-weighted portfolio and computing its effective number of assets, $N_{\text{eff},t}^{\text{agg}}$ is not a benchmark for an individual investor. Aggregating heterogeneous and concentrated portfolios mechanically raises N_{eff}, so $N_{\text{eff},t}^{\text{agg}}$ will exceed the mean individual effective number $\bar{N}_{\text{eff},t}$ in any market with heterogeneous investors; i.e., can increase the effective number of holdings through differences in portfolio composition, although it does not necessarily produce a value exceeding the arithmetic mean of individual effective numbers. In addition, an individual investor cannot costlessly replicate the aggregate portfolio when there are fixed costs per holding.

The aggregate effective number of holdings describes the concentration of the portfolio obtained by combining all observed client holdings at date t. The number of distinct financial instruments held by at least one client at that date, M_t), describes the breadth of the observed platform holdings. Both quantities demonstrate financial instruments and portfolio compositions observed within the client base, but neither establishes the effective opportunity set available to every individual investor.

The distinction between date-specific and pooled counts is important. Table 18 reports

distinct financial instruments held at each observation date, whereas Table 17 counts trading symbols held at least once across all four dates. The pooled count should not be denoted by M_t, which refers to a single date.

Hypothesis 4 (observed investment opportunities and individual portfolio breadth). The number of distinct financial instruments held across the broker's client base substantially exceeds the number held by individual investors. The aggregate effective number of holdings is also compared with the distribution of individual effective numbers. These comparisons assess the observed breadth of platform holdings relative to individual portfolio choices, without establishing that all investors face identical opportunity sets or that individual access constraints are nonbinding.

H4 differs from H1–H3 because it is not a restriction on the parameters of an estimated equation. We assess it by comparing the attainability bound $N_{\text{eff},t}^{\text{agg}}$ with the empirical distribution of $N_{\text{eff},it}$ across the investors observed at the same date. We report where the aggregate effective number falls in the individual distribution and how many investors reach it. These comparisons describe observed portfolio concentration and do not constitute a statistical rejection of limited investment opportunities as an explanation for narrow individual portfolios.

One qualification applies to M_t. Because the broker offers international access and markets itself partly on that basis, its clientele may be selected on wanting such access. The resulting M_t may therefore overstate the opportunity set facing a Romanian household at an arbitrary intermediary. The aggregate effective number of holdings $N_{\text{eff},t}^{\text{agg}}$ is computed from the same client population and is therefore also conditional on selection into this broker. The observed count of domestic financial instruments is not affected by the availability of foreign markets, but it remains distinct from the set of domestic instruments practically accessible to every individual investor.

A final element of the framework is an accounting identity that guides the interpretation of time-series comparisons. When the investor population grows rapidly and entrants differ systematically from incumbents, changes in cross-sectional averages need not reflect changes in behavior.

Partition investors into stayers S (present at both dates), entrants E (present only at $t+1$), and exiters X (present only at t). For any investor-level statistic b, write $\bar{b}_t$ for its cross-sectional mean at date t. Then

$$\underbrace{\bar{b}_{t+1} - \bar{b}_t}_{\text{total}} = \underbrace{\left(\bar{b}_{S,t+1} - \bar{b}_{S,t}\right)}_{\text{within stayers}} + \underbrace{\frac{|E|}{N_{t+1}}\left(\bar{b}_{E,t+1} - \bar{b}_{S,t+1}\right)}_{\text{entry}} - \underbrace{\frac{|X|}{N_t}\left(\bar{b}_{X,t} - \bar{b}_{S,t}\right)}_{\text{exit}}, \tag{8}$$

which holds exactly (Appendix A). Here, n_t denotes the number of investors observed at date t, to distinguish the sample size from individual portfolio breadth N_{it}.

Table 1 summarises the hypotheses, the corresponding estimating relationships, and the associated null and alternative hypotheses.

Table 1 Hypotheses, estimating equations and tests

	Hypothesis	Equation	Null and alternative	Reported in
H1a	Breadth increases with portfolio value, log-linearly	(2), (5)	$H_0 : \beta = 0$; $H_1 : \beta > 0$	§5.2
H1b	Demographic gradients operate through portfolio value	(3), (4)	$H_0 : \tilde{\gamma}_k = \gamma_k$; $H_1 : \lvert\tilde{\gamma}_k\rvert < \lvert\gamma_k\rvert$	§5.2
H2	Delegation intensity falls and participation rises with portfolio value	(6)	$H_0 : \beta_f = \beta_p = 0$ (Wald); intersection–union test of $H_1 : \beta_f < 0$ and $\beta_p > 0$	§5.3
H3	The domestic share responds less than breadth to portfolio value, within investor	(7)	$H_0 : \eta_N - \lvert\eta_s\rvert = 0$; $H_1 : \eta_N - \lvert\eta_s\rvert > 0$	§5.4
H4	The opportunity set does not bind	—	Not a parameter restriction; $N_{\text{eff},t}^{\text{agg}}$ compared with the distribution of $N_{\text{eff},it}$	§5.5

3.3 Identification and inference

The hypotheses describe conditional relationships in the data rather than causal treatment effects. For example, H1a states that investors with larger portfolios tend to hold more distinct financial instruments in this population. It does not claim that an exogenous increase in portfolio value would necessarily cause an investor to add holdings. The empirical assessments are designed to evaluate observable implications of portfolio-choice theories, recognizing that several mechanisms can generate similar conditional relationships.

Portfolio value is jointly determined with portfolio composition. A concentrated position that appreciates mechanically raises portfolio value while also raising concentration, and changes in portfolio value may therefore be endogenous to changes in portfolio breadth and concentration. The within-investor specifications in (5) and the second equation of (7) remove time-invariant investor heterogeneity, but they do not resolve this endogeneity. Section 5.2 additionally reports an exploratory instrumental-variable (IV) specification using passive portfolio returns. Although this approach isolates a component of portfolio-value variation associated with changes in the prices of previously held financial instruments, its validity as a causal identification strategy depends on an exclusion restriction that the available data do not establish. The IV estimates are therefore interpreted as a sensitivity exercise rather than as identified causal effects.

The tests of H1–H4 therefore assess selected empirical implications of household-finance theories and describe the range of financial instruments and portfolio compositions observed in the data, without identifying structural parameters. For this reason, we avoid causal language in Section 5. We also distinguish statistical support for a directional association from evidence that a particular theoretical mechanism generated it.

We report standard errors clustered at the investor level. The data are a census of one broker's active clients rather than a random draw from Romanian households, so classical sampling uncertainty is not the natural concept for the population observed. Reporting clustered standard errors is nevertheless useful as a conventional description of precision for this type of client base.

The empirical focus is on economic magnitude and precision, not only on whether coefficients differ from zero. In particular, failure to reject a null is not interpreted as evidence in favour of the null. When a coefficient is imprecisely estimated, we report the point estimate and its uncertainty without drawing a binary conclusion.

Because the sample contains between roughly 35,000 and 96,000 investor-years across the reported empirical specifications, conventional statistical significance is often inexpensive. Economic magnitude is therefore emphasized throughout.

The four hypotheses also differ in the sharpness of their empirical content. H1 and H2 are examined using conditional regression relationships, with directional coefficient restrictions for H1a and H2; H3 is a restriction on a function of estimated coefficients, and H4 is a descriptive comparison of aggregate and individual portfolio characteristics. H4 carries no sampling distribution in the empirical framework presently specified. It can establish that the observed platform holdings extend beyond the portfolios of individual clients, but it does not identify what an individual optimum would be or demonstrate that every investor faces an unconstrained opportunity set. For this reason, statistical and economic significance are reported separately.

4 Data

4.1 Source, sample and cleaning

The data are a proprietary, anonymized extract from a large Romanian brokerage firm. They cover every position held by every active retail client at four observation dates: 31 December 2020, 31 December 2021, 31 December 2022, and 22 August 2023. Each position includes a pseudonymized investor identifier, the trading symbol, quantity held, last observed market price, and the trading currency. Positions are valued at market and converted to Romanian leu using the official exchange rate of the National Bank of Romania.[3]

[3] At the end of the sample (22 August 2023), the National Bank of Romania reference exchange rate was 1 EUR = 4.9373 RON and 1 USD = 4.5276 RON (equivalently, 1 RON = 0.2025 EUR and 0.2209 USD).

A financial instrument is operationally identified by its trading symbol in the brokerage records. The dataset therefore distinguishes the number of recorded positions from the number of distinct trading symbols held. Market-wide statistics on listed companies and issuers, discussed in Section 2, describe different objects and are not used interchangeably with these instrument-level counts.

A separate file records background characteristics collected at account opening. These include age, sex, and place of birth and residence as documented on the Romanian identity card. Education, income band, and prior investment experience are self-reported, optional, and not externally validated. A third file maps trading symbols to financial instrument type, issuer country, and sector. The extract is a census of one broker's active clients. Each observation is an end-of-day snapshot of holdings rather than a record of transactions. Trades, their timing, and turnover are not observed. Consequently, changes in observed positions between two dates cannot generally be attributed to purchases, sales, transfers, or price changes separately. Similarly, the first appearance of an investor in the observed sample does not necessarily identify the date on which the investor opened a brokerage account or first participated in financial markets.

All portfolio values and other monetary quantities derived from the holdings data are in current Romanian leu at the observation date and are not deflated. Inflation was high over this period, so nominal changes in portfolio value differ from changes measured in constant purchasing-power terms. This affects statements about levels and growth in portfolio value.

Inflation does not affect the portfolio-breadth and concentration measures reported in this paper when all position values at a given observation date are converted using the same price deflator. The number of distinct financial instruments, the Herfindahl–Hirschman index, the effective number of holdings, and all value shares are unchanged by any within-date price deflator. Cross-sectional gradients in $\ln V_{it}$ are also unaffected by expressing portfolio values in constant prices because deflation adds a date-specific constant that is absorbed by date fixed effects. This invariance does not extend automatically to the interpretation of nominal portfolio-value changes across observation dates, particularly when changes are related to changes in investor behavior.

The four observation dates are not equally spaced, and the last date is not a year-end snapshot. The first two intervals span 365 days, while the last runs from 31 December 2022 to 22 August 2023, a span of 234 days. The final interval is therefore right-censored. Entry and exit occurring between 22 August 2023 and 31 December 2023 are not observed, and observed transitions in investor participation and portfolio composition cover approximately 7.7 months rather than a full year. Section 5.6 adjusts entry and exit rates for comparability. Any statistic computed on the 2023 cross-section describes an August position rather than a year-end position. Section 5.6 also repeats the analysis on the three year-end dates alone.

We drop empty positions by excluding holding lines with market value below RON 0.005. This removes zero-price records and negligible residual lines and yields 766,637 positions.

Investor-level statistics are computed on 9,837, 18,206, 26,943, and 41,234 investor-years, for a pooled total of 96,220 investor-years and 43,360 distinct investors.

The unit of observation is an investor at an observation date. We use the conventional term investor-year, noting that the fourth observation is in August rather than December. The composition tables in Appendix C and the distributions in Appendix D were independently recomputed from the raw extract and reproduce to published precision.

One classification caveat remains. For 1,226 trading symbols, the country classification is unavailable in the manually assigned field and is therefore taken from a less detailed country field. These symbols account for 1.67% of held value. Table 13 reports that the maximum effect of this uncertainty on the value-weighted domestic portfolio share is at most 0.13 percentage points in any year.

Table 2 summarizes the sample. Aggregate brokerage portfolio value grows from RON 2.15 billion to RON 5.17 billion, while the median investor portfolio value falls from RON 26,424 to RON 15,524. This reflects a participation wave in which new entrants are much smaller than incumbents. Portfolio value is also highly unequal, with a Gini coefficient of 0.85 and a top-percentile share of 41% in 2023, although inequality falls slightly as the client base broadens.

Table 2 Sample overview by observation date

	2020	2021	2022	2023
Investors	9,837	18,206	26,943	41,234
Positions	68,259	153,009	225,018	320,351
Total value (mln RON)	2,152.05	3,392.06	3,723.66	5,169.96
Mean portfolio (RON)	218,770.65	186,315.71	138,205.00	125,380.96
Median portfolio (RON)	26,424.20	21,494.22	17,590	15,524.04
P90 portfolio (RON)	314,940.36	288,187.68	222,138.86	215,076.69
P99 portfolio (RON)	2,558,372.42	2,455,626.36	1,673,085.34	1,531,108.43
Top 1% share of value (%)	48.12	44.69	42.88	41.19
Top 10% share of value (%)	79.53	79.12	77.82	77.28
Gini of portfolio value	0.87	0.86	0.85	0.85

Notes: Positions with market value below RON 0.005 excluded. "Positions" counts non-empty holding lines. "Investors" counts distinct investor identifiers observed at each date. Portfolio values are reported in RON using observation-date National Bank of Romania exchange rates. The Gini coefficient is computed across investors within each observation date. The August 2023 date is not a year-end observation, so cross-year comparisons involve unequal observation intervals.

4.2 Investor characteristics, the extreme tail and representativeness

Self-reported background characteristics are missing for a large share of clients, and the missingness follows a small number of structured patterns. Only 8 of the 32 logically possible missingness patterns occur. Just over half of investor-years (51.8%) report all characteristics.

A single pattern in which education, income, and experience are all missing accounts for a further 32.0%. Because these missingness indicators are highly collinear, estimates that include multiple non-disclosure dummies can be sensitive to specification. We therefore estimate each demographic relationship on the subsample that reports the characteristic in question. Non-disclosure is analyzed directly in Section 5.6. The resulting estimates describe relationships among investors for whom the relevant characteristic is observed.

The term *retail* is the broker's classification for a non-institutional account. It is not a statement about wealth. The distribution of account values is extremely skewed. In 2023, the largest account holds RON 370.9 million, which is 7.2% of the sample's total value. The five largest accounts hold 9.8%. The 31 accounts above RON 10 million hold 18.7%, and the 733 accounts above RON 1 million hold 48.7%. Concentration is also visible at the security level. Of the 46 securities with more than RON 20 million held across the broker's client base in 2023, one security, the largest government bond line, has 97.2% of its value held by sampled clients in a single account.

Whether to impose a retail wealth cut-off is a research-design decision and is not taken here. We use the full sample for the main analysis. Table A7 reports sensitivity to trimming the top 0.1%, 1%, and 5% of accounts by value within each observation date. Dropping the top 1% removes 43.5% of aggregate value but moves mean breadth from 7.97 to 7.88 holdings, mean N_{eff} from 4.25 to 4.23, and the value-weighted domestic share from 86.9% to 86.5%.

The sample is not representative of Romanian households in two clear ways. First, it is conditional on participation, and participants differ from non-participants in wealth, education, financial literacy, and trust (Haliassos and Bertaut, 1995; van Rooij et al., 2011; Guiso et al., 2008). Second, it is conditional on this broker, which is a full-service domestic intermediary offering both domestic and international access, no leveraged products, and Romanian-language support. These features may influence the composition of the broker's clientele, but the available data do not establish how its clients differ from those using other intermediaries. The empirical results should therefore be interpreted as describing the observed broker population rather than Romanian retail investors generally.

Relative to Bucharest Stock Exchange aggregates (Appendix Table A11), the broker's clients hold a small share of domestic equity market capitalization, ranging from 0.61% to 0.91% in 2022, before declining to 0.81% in in 2023. Their share of the reported domestic bond-market capitalization is substantially larger, ranging from 10.7% to 14.1% over the four observation dates. These figures provide an indication of the broker's importance in the observed market segments, particularly in government bonds. However, the 2023 comparisons are not strictly contemporaneous because the brokerage holdings are measured on August 22, whereas the exchange statistics refer to year-end.

5 Results

5.1 Portfolio breadth and concentration

Table 3 reports the distribution of the number of distinct financial instruments held and several measures of portfolio concentration by observation date. Appendix D reports the full percentile distributions. The number of distinct financial instruments measures portfolio breadth, whereas the Herfindahl–Hirschman index and its reciprocal describe the concentration of portfolio value across those instruments. These measures do not incorporate correlations among asset returns or the underlying exposures of investment funds and ETFs.

Pooling investor-date observations, the mean investor holds 7.97 distinct financial instruments and the median holds five. The distribution is highly skewed. Portfolios containing a single financial instrument account for 18.1% of investor-date observations, 36.9% hold three distinct financial instruments or fewer, and 8.0% hold 20 or more. Classical studies of equity portfolio diversification examine how increasing the number of stocks reduces idiosyncratic risk (Evans and Archer, 1968; Statman, 1987). However, the number of financial instruments held in the present data cannot be interpreted using a universal threshold, since the portfolios include equities, bonds, funds, and other financial instruments. Figure 2 shows a large mass at one holding and a long tail reaching 200 distinct financial instruments. Appendix Table A18 shows that right-skewed distribution persists across observation dates.

Simple counts can overstate diversification when portfolios are concentrated in a few large positions. The mean Herfindahl–Hirschman index is 0.452 and the median is 0.334 (Appendix Table A19). The mean effective number of assets is 4.25 and the median is 3.00 (Appendix Table A20). Thus, the median effective number of holdings is approximately three, compared with a median portfolio breadth of five distinct financial instruments. These two statistics describe separate marginal distributions and do not necessarily characterize the same investor. Their difference indicates that portfolio breadth alone does not capture the concentration of value across financial instruments.

Two additional measures are reported for completeness but are less informative for the principal empirical questions examined here. The Gini coefficient of portfolio weights is zero both for single-asset and for perfectly equally weighted portfolios, which is why its lower percentiles are at zero (Appendix Table A21). Normalized Shannon entropy averages 0.776 among portfolios containing at least two distinct financial instruments, but by construction it measures balance conditional on the number of holdings. It is therefore not a direct measure of diversification and is undefined for single-asset portfolios (Appendix Table A22).

Portfolio breadth and concentration changes over time in a non-monotone way. Mean breadth rises from 6.94 in 2020 to 8.40 in 2021, is 8.35 in 2022, and falls to 7.77 in 2023. The share of portfolios containing a single financial instrument falls from 21.8% to 16.6% in 2022 and then rise to 18.8% in August 2023. Section 5.6 shows that the 2023 reversal is

largely attributable to the composition of the observed client base, particularly the entry of investors with narrower portfolios.

Narrow portfolios account for a large share of investors but a smaller share of portfolio value. In 2023, portfolios containing one financial instrument account for 18.8% of investors but only 6.9% of the aggregate brokerage portfolio value. Portfolios with $N_{\text{eff}} < 2$ account for 38.3% of investors and 30.9% of value, while in the pooled sample, 7.1% of investor-date observations containing more than 20 distinct financial instruments account for 20.4% of aggregate portfolio value (Appendix Table A3).

The value-weighted mean number of distinct financial instruments in 2023 is 14.66, compared with an equal-weighted mean of 7.77 (Table 4). The equal-weighted statistic describes the average portfolio breadth across investors, while the value-weighted statistic describes the average breadth of investors' portfolios when each investor is weighted by the value of their observed brokerage holdings. The corresponding effective numbers of holdings are 5.74 and 4.11, respectively. Thus, larger brokerage portfolios receive greater weight in the value-weighted statistics and are associated, on average, with both greater portfolio breadth and a higher effective number of holdings.

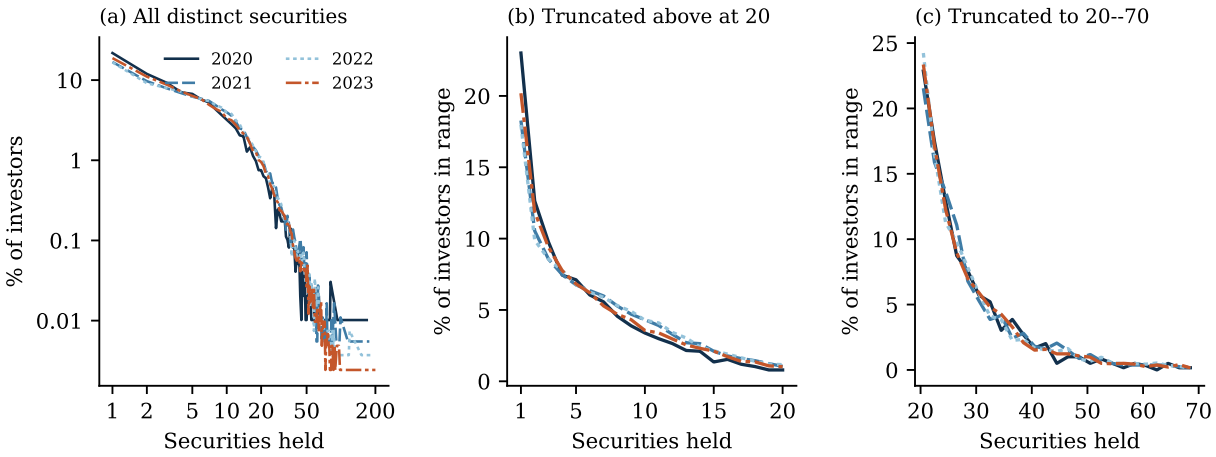


Figure 2 Distribution of portfolio breadth

Notes: Each line is an observation date. Panel (a) shows the whole support on logarithmic axes; the maximum observed is 200 instruments. Panels (b) and (c) truncate the sample rather than top-code it: panel (b) retains investor-years with $N_{it} \leq 20$, which is 92.9% of the pooled sample, and panel (c) retains those with $20 \leq N_{it} \leq 70$, which is 97.5% of the investor-years above the lower truncation point. No observation is moved, and percentages in panels (b) and (c) are of the investor-years within the respective range. Both axes in panel (a) are logarithmic, with decimal tick labels.

5.2 Portfolio value, breadth and investor characteristics

Table 8 sorts investors into deciles of brokerage portfolio value within each observation date. Median portfolio value rises from RON 324 in the first decile to RON 486,227 in the tenth, a factor of about 1,500. Over this range, the mean number of distinct financial instruments held rises from 2.46 to 15.24, and the mean effective number of holdings rises from 1.81 to

Table 3 Breadth and concentration by observation date

	2020	2021	2022	2023	All
N	9,837	18,206	26,943	41,234	96,220
Mean N	6.94	8.40	8.35	7.77	7.97
Median N	5	6	6	5	5
P90 N	16	19	18	18	18
Share $N = 1$ (%)	21.76	16.82	16.64	18.80	18.12
Share N$\leq$3 (%)	42.95	34.40	33.72	38.61	36.89
Share N$\geq$20 (%)	6.15	8.94	8.50	7.80	8.04
Mean HHI	0.50	0.43	0.43	0.46	0.45
Median HHI	0.40	0.31	0.30	0.35	0.33
Mean N_{eff}	3.76	4.50	4.46	4.11	4.25
Median N_{eff}	2.49	3.28	3.32	2.89	3.00
Share $N_{\text{eff}}<2$ (%)	43.68	34.03	33.62	38.32	36.74
Value in $N_{\text{eff}}<2$ (%)	49.05	36.58	33.64	30.90	35.65
Value in $N = 1$ (%)	11.76	7.54	7.92	6.88	8.03

Notes: N_{eff} as defined in (1). "Value in $N_{\text{eff}} < 2$" is the share of total portfolio value held by investors whose effective number of assets is below two. Full percentile distributions are in Appendix D.

Table 4 Equal-weighted and value-weighted diversification

	2020	2021	2022	2023
EW mean N	6.939	8.404	8.352	7.770
VW mean N	9.565	13.572	14.026	14.658
EW mean N_{eff}	3.763	4.501	4.456	4.112
VW mean N_{eff}	4.134	5.460	5.619	5.738
EW mean HHI	0.501	0.432	0.430	0.463
VW mean HHI	0.533	0.424	0.415	0.390

Notes: EW denotes the equally weighted mean across investors, and VW denotes the mean weighted by the value of each investor's brokerage portfolio. Both statistics average investor-level measures; neither represents the effective number of holdings in the aggregate client portfolio.

6.27. The share of portfolios containing a single financial instrument falls from 50.6% to 6.6%, and the mean Herfindahl–Hirschman index falls from 0.748 to 0.333. Participation in foreign financial instruments rises from 8.0% to 47.2%, and participation in government bonds rises from 1.0% to 41.9%. Each of these profiles is monotone.

Estimating Equation (2) on the pooled panel yields $\hat{\beta} = 1.593$ (clustered standard error 0.024). In economic terms, a doubling of portfolio value is associated with 1.10 additional instruments held, and a tenfold increase with 3.67 (Table 9). When $\log N_{\text{eff}}$ is used as the outcome, the coefficient is 0.135 (standard error 0.002), implying 0.094 increase in log effective holdings, equivalent to approximately a 9.8% increase in the effective number of holdings for a doubling of portfolio value. Adding the full set of account-opening characteristics and non-disclosure indicators changes the two coefficients only slightly, to 1.497 and 0.133.

An approximately log-linear specification summarizes the relationship between portfolio breadth and portfolio value reasonably well over the observed distribution. Applying $\hat{\beta}$ to the 1500-fold value range between the first and tenth deciles predicts 11.6 additional instruments, compared with 12.8 observed. The left panel of Figure 3 shows some departures from a strictly linear relationship in log portfolio value.

The evidence also suggests that larger portfolios contain more distinct financial instruments, while their portfolio value remains relatively concentrated. From the first to the tenth value decile, the raw number of instruments held rises by a factor of 6.2, while the effective number of holdings rises by a factor of 3.5. As a result, the ratio N_{eff}/N calculated from the decile means falls from 0.74 to 0.41.

The cross-sectional portfolio-value gradient could partly reflect sorting across investor types. Table 7 reports within-investor estimates from Equation (5). Among 26,999 investors observed at least twice, the within-investor coefficient is 1.780 (standard error 0.043). In the balanced panel of 8,718 investors observed at all four dates, it is 1.785 (standard error 0.068). Both exceed the cross-sectional estimate of 1.593. These results indicate that the positive association between portfolio value and breadth also appears within investors over time. However, the cross-sectional and within-investor coefficients rely on different sources of variation and do not necessarily estimate the same economic relationship.

This comparison does not rule out sorting, because the within-investor estimator has a mechanical component. Portfolio value is the sum of position values, acquiring a previously unheld financial instrument increases portfolio breadth, while its effect on total portfolio value depends on how the purchase is financed. The component of within-investor variation in $\ln V$ that is attributable to newly opened positions accounts for 18.1%, 12.3%, and 32.3% of the absolute variation in the three transitions. These accounting contributions should not be interpreted as direct measures of net deposits or active portfolio expansion, since an investor may finance a new position by selling an existing financial instrument.

Table 5 addresses this issue by using passive portfolio returns. Following Calvet et al.

(2009), passive value is constructed as the value at date t of the instruments held at $t-1$, revalued at date-t prices. Changes in passive value therefore reflect the hypothetical change in the value of previously observed holdings arising from their revaluation, rather than the addition of newly held financial instruments. We estimate the first-difference specification over consecutive observation dates,

$$\Delta N_{it} \;=\; \alpha_t \;+\; \beta^{FD}\,\Delta \ln V_{it} \;+\; \varepsilon_{it}, \tag{9}$$

on 48,103 investor-pairs covering 25,919 investors whose start-of-window portfolio is at least 90% priced.

We additionally use the passive portfolio return as an instrumental variable (IV) for the observed change in log portfolio value. This specification is intended to assess whether the positive within-investor association is also present in portfolio-value variation associated with revaluations of previously held financial instruments. It is not treated as a causal identification strategy.

Ordinary least squares yields a coefficient of 1.7 (standard error 0.036), close to the fixed-effects estimate. The passive-return IV is strongly associated with the observed change in log portfolio value in the first-stage regression, whose estimated coefficient is 0.230 (standard error 0.015). However, the 2SLS estimate is estimate is -0.234 (standard error 0.164), so the hypothesis of no within-investor response cannot be rejected at conventional levels. The 2SLS estimate therefore differs substantially from the positive OLS association and is not statistically distinguishable from zero. Because the exclusion restriction is not established, this difference cannot be interpreted as identifying the causal effect of passive changes in portfolio value on portfolio breadth. The final observation interval is also shorter than one year.

Two interpretations remain. The passive-return IV is based on revaluations of previously held financial instruments, whereas the fixed-cost mechanism is about the relationship between the benefits and costs of expanding portfolio breadth, which may differ between passive valuation changes and additional investable resources. A weak one-year response is therefore potentially consistent with adjustment costs, portfolio inertia, or differences in how investors respond to valuation changes and additional funds (Bilias et al., 2010). Alternatively, the cross-sectional gradient may reflect sorting, with investors who plan to hold many securities also being those who fund their accounts. We therefore report the portfolio-value gradient as a cross-sectional fact, without treating the within-investor estimates as evidence against sorting. These mechanisms are not mutually exclusive, and the available holdings snapshots do not permit us to determine their relative contributions.

A simpler explanation for the portfolio-value gradient is a minimum practical holding size. If each security must be purchased in a non-trivial amount, small portfolios may be unable to hold many instruments regardless of preferences. Table 6 able 6 examines the

relationship between portfolio value, the typical size of recorded positions, and the number of distinct financial instruments held. Although this comparison provides descriptive evidence on portfolio scale, it does not directly establish the existence or absence of binding minimum investment requirements.

Holding sizes rise with portfolio value. The median holding increases from RON 82 in the first value decile to RON 12,369 in the tenth. The ratio of median portfolio value to median recorded position size, computed as the decile's median portfolio value divided by the median holding size in the same decile, rises from 3.9 to 39.3. Over the same range, the mean number of instruments held rises from 2.46 to 15.24. Mean portfolio breadth represents between 39% and 63% of this descriptive ratio across the portfolio-value deciles, with lower values in the larger-portfolio deciles. However, the ratio is not a measure of the maximum number of financial instruments an investor could afford to hold, because the median observed position size is not an exogenous minimum investment requirement.

Functional form points in the same direction. Under additional assumptions, a binding minimum investment amount for each position could generate a relationship between portfolio breadth and portfolio value. Models incorporating fixed information or transaction costs can generate other relationships, depending on the assumed cost structure and investor preferences.

In pooled data, the log specification fits substantially better ($R^2 = 0.174$) than the level specification ($R^2 = 0.001$). In the bottom half of the value distribution the two fits are similar (0.100 versus 0.097), so minimum holding sizes may still matter at the very bottom. These comparisons indicate differences in statistical fit but do not identify the economic mechanism underlying the portfolio-value gradient.

Table 5 The within-investor gradient with a passive-value instrument

Specification	Coefficient	SE	Pairs
OLS, first differences	1.700***	(0.036)	48,103
First stage: $\Delta \ln V$ on passive return	0.230***	(0.015)	48,103
Reduced form: ΔN on passive return	-0.054	(0.036)	48,103
2SLS, passive return as instrument	-0.234	(0.164)	48,103

Notes: Equation (9) in first differences over consecutive observation dates, 48,103 investor-pairs and 25,919 investors, restricted to investors whose start-of-window portfolio is at least 90% priced; median coverage is 100%. The instrumental variable (IV) is the passive log return calculated from financial instruments held at the beginning of each observation interval, defined as the value at t of the instruments held at $t-1$ revalued at t prices, divided by their value at $t-1$. The first-stage relationship establishes relevance of the passive-return IV but does not establish the exclusion restriction required for causal interpretation. The 2SLS results are therefore reported as a sensitivity analysis. The final observation interval is shorter than the preceding annual intervals. All specifications include transition fixed effects, and standard errors are clustered by investor. The passive return removes the arithmetic link between adding a position and raising portfolio value, but it is not exogenous in a structural sense: returns on concentrated holdings are correlated with the type of investor who holds them. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 10 examines H1b by estimating Equations (3) and (4) on the subsample that

Table 6 Portfolio value, typical position size, and portfolio breadth

	1	2	3	4	5	6	7	8	9	10
Median value (RON)	324.40	2,253.31	4,689.82	8,202.67	13,664.74	23,097.44	39,886.71	72,351.93	148,633.14	486,226.91
Mean N	2.46	4.00	5.04	6.01	7.02	8.10	9.31	10.40	12.10	15.24
Median position (RON)	82.44	323.40	504.40	717.12	1,000.80	1,365	1,962.95	2,971.50	4,986.17	12,368.83
Implied capacity	3.93	6.97	9.30	11.44	13.65	16.92	20.32	24.35	29.81	39.31
Mean N / capacity	0.63	0.57	0.54	0.53	0.51	0.48	0.46	0.43	0.41	0.39

Notes: Implied capacity is the decile's median portfolio value divided by the median position size of investors in the same decile, i.e. the number of typical positions that portfolio could support. The reported ratio compares median portfolio value with median recorded position value within each decile. It is a descriptive measure of relative portfolio and position scale, not an estimate of the maximum number of financial instruments that an investor can afford to hold. Functional-form comparisons are in Appendix Table A5.

reports each characteristic. The raw comparisons reproduce familiar patterns. Women hold 1.90 fewer securities than men. Investors aged 18–24 hold 1.87 fewer securities than the 25–34 benchmark, while those aged 35–44 hold 1.23 more. Investors with only high school education hold 0.96 fewer than university graduates, and those with post-secondary non-tertiary education hold 1.74 fewer. Income shows a generally increasing gradient across the reported income bands up to EUR 120,000, from +0.73 for the 6–12k band to +2.84 for the 48–120k band. These comparisons are conditional on observation date but do not control for portfolio value or other investor characteristics. They describe differences relative to the relevant reference category within the sample reporting each characteristic.

Adding $\ln V_{it}$ substantially reduces several of these estimated differences, although the magnitude and direction of the changes vary across categories. For age, the 18–24 deficit falls from −1.87 to −0.25 and the 35–44 advantage from +1.23 to +0.05, indicating that the corresponding differences in portfolio breadth are much smaller among investors with similar brokerage portfolio values (Roche et al., 2013). However, meaningful conditional differences remain in several categories, including older investors, those with post-secondary non-tertiary education, and those reporting income above EUR 120,000. Measured by the share absorbed proportional attenuation of the corresponding unadjusted coefficient, the estimates decline by 87% for the 18–24 deficit, 96% for the 35–44 advantage, and 62–88% across the income gradient between EUR 6,000 and EUR 120,000.

Conditioning on brokerage portfolio value substantially attenuates several demographic and socioeconomic differences in portfolio breadth, particularly among younger investors and middle-income groups. However, sizable conditional differences remain for certain age, education, and high-income categories, indicating that the cross-sectional heterogeneity is not summarized by portfolio value alone. For sex, the pattern differs. The female deficit falls only from −1.90 to −1.42, he conditional difference is approximately 25% smaller than the unadjusted difference. Women therefore hold about 1.4 fewer securities than men with portfolios of similar size in the observed sample. However, the conditional regression in Table 10 controls for portfolio value and observation date, not necessarily for all other demographic and socioeconomic characteristics simultaneously. Thus, the estimated gender difference should not be interpreted as a comparison between otherwise identical men and women. Differences in age, education, income, investment experience, or other unobserved characteristics could remain relevant.

Two other conditional patterns reverse sign. For investors aged 65 and over, the coefficient moves from −0.46 in the raw comparison to −2.53 conditional on portfolio value. For the top income band, it moves from +2.04 to −2.12. Appendix C shows that both groups are small in number but hold a large share of the observed brokerage portfolio value. In 2023, investors aged 65 and over are 3.5% of the client base and hold 16.5% of value (Appendix Table A13). Those reporting income above EUR 120,000 are 1.6% of investors and hold 23.0% of value (Appendix Table A15). Conditional on portfolio value, both groups

hold relatively few distinct financial instruments compared with their respective reference categories. These patterns may reflect differences in portfolio preferences, investment objectives, financial circumstances, or unobserved holdings outside the brokerage platform. However, the available data do not identify which of these explanations is responsible.

The composition tables also place the gender result in context. Women are 24.7% of the 2023 client base but hold 16.1% of its observed brokerage portfolio value (Appendix Table A12). Appendix Tables A14 and A16 show corresponding concentration patterns for education and reported investment experience: university and postgraduate investors are 43.7% of the client base and hold 75.2% of value, and investors reporting more than three years of experience are 4.6% of investors and 21.8% of value. These figures describe the composition of the observed brokerage client base. They should not be interpreted as estimates of the distribution of financial wealth, educational attainment, or investment experience among Romanian households. Moreover, the education and experience comparisons are affected by non-disclosure of these characteristics.

Reported investment experience is the least interpretable characteristic. Relative to the less-than-six-months benchmark, the raw differences are small and negative (-0.61 for 6–12 months and -0.49 for more than three years), while the conditional coefficient for the most experienced group is -2.14. The estimates indicate that investors reporting more than three years of prior investment experience hold fewer distinct financial instruments than the reference group after conditioning on brokerage portfolio value. This association does not establish whether greater experience leads investors to choose more concentrated portfolios, whether investors with different portfolio preferences acquire different levels of experience, or whether other characteristics account for the relationship.

However, this variable is measured only once at account opening. The data do not record how long before each observation date the account was opened, the measure is missing for 61% of clients, and most reporters describe themselves as novices. We therefore report these gradients without assigning them a causal or behavioral interpretation.

This limitation concerns measurement rather than the underlying question. Campbell et al. (2014) show that investment experience shapes portfolio behavior. The dates at which investors first appear in our panel allow us to measure the duration of their observed participation in the sample, but not their actual account tenure or total investment experience. We leave a more detailed analysis of investment experience and portfolio adjustment for future work.

Taken together, these results show that brokerage portfolio value is strongly and positively associated with portfolio breadth, both across investors and within investors over time. Conditioning on portfolio value substantially reduces several demographic and socioeconomic differences in breadth, although meaningful heterogeneity remains. These patterns are consistent with the potential importance of financial capacity and investor-specific portfolio-choice frictions. However, the available evidence does not identify the relative con-

Table 7 Within-investor estimates

Sample	Outcome	b(ln value)	SE	Investor-years	Investors
Investors seen ≥ 2 times	Portfolio breadth (N)	1.780***	(0.043)	79,859	26,999
Investors seen ≥ 2 times	Domestic share	-0.005***	(0.001)	79,859	26,999
Investors seen ≥ 2 times	Fund/ETF value share	-0.008***	(0.001)	79,859	26,999
Balanced panel (all four)	Portfolio breadth (N)	1.785***	(0.068)	34,872	8,718
Balanced panel (all four)	Domestic share	-0.003	(0.002)	34,872	8,718
Balanced panel (all four)	Fund/ETF value share	0.001	(0.002)	34,872	8,718

Notes: Investor and observation date two-way within estimator (5) and the second equation of (7); the coefficient is on $\ln V_{it}$. Standard errors clustered by investor, with degrees of freedom corrected for the absorbed investor effects. Coefficients are estimated using within-investor variation after removing investor and observation-date fixed effects, so every time-invariant investor characteristic is differenced out, but $\ln V_{it}$ remains jointly determined with composition. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

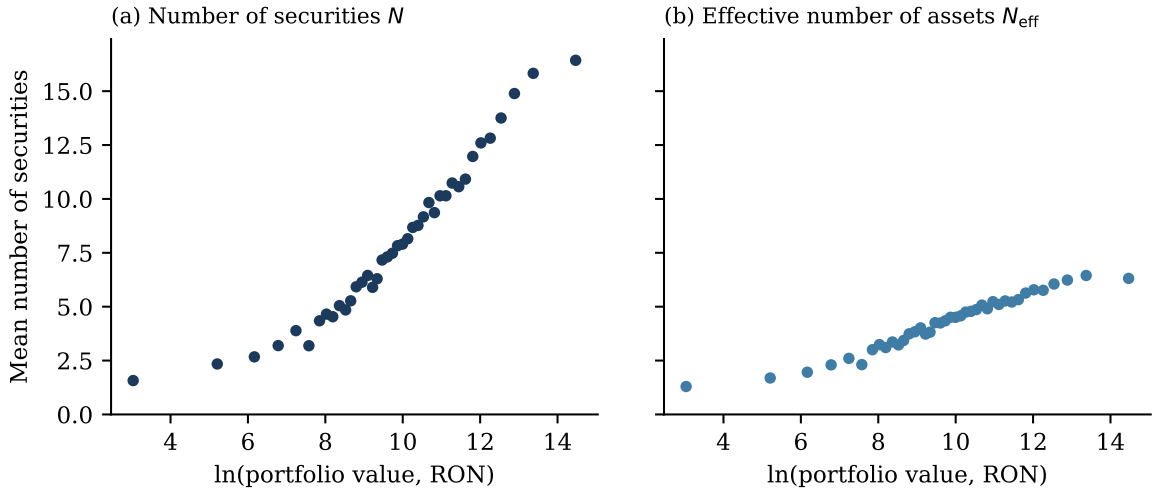


Figure 3 Portfolio value, portfolio breadth, and the effective number of holdings

Notes: The figure compares conditional means of portfolio breadth and the effective number of holdings over the pooled investor-date observations. Their divergence at higher portfolio values indicates that increases in the number of distinct financial instruments are not matched proportionally by increases in the concentration-equivalent number of holdings.

Table 8 Portfolio structure by decile of portfolio value

	1	2	3	4	5	6	7	8	9	10
Median value (RON)	324.4	2,253.3	4,689.8	8,202.7	13,664.7	23,097.4	39,886.7	72,351.9	148,633.1	486,226.9
Mean N	2.5	4.0	5.0	6.0	7.0	8.1	9.3	10.4	12.1	15.2
Median N	1	3	4	5	6	7	8	9	10	12
Mean N_{eff}	1.8	2.8	3.4	3.8	4.2	4.5	4.9	5.2	5.6	6.3
Mean HHI	0.7	0.6	0.5	0.5	0.4	0.4	0.4	0.4	0.3	0.3
Single-asset (%)	50.6	28.6	22.5	17.1	14.5	11.8	11.2	10.1	8.2	6.6
Holds foreign asset (%)	8.0	14.7	18.5	22.1	26.6	32.2	35.5	37.9	41.6	47.2
Domestic (%)	94.8	93.7	93.7	93.4	92.5	90.9	90.4	89.7	88.5	85.6
Holds fund/ETF (%)	26.0	41.3	46.1	49.2	51.9	53.2	53.8	53.2	51.2	49.9
Fund/ETF value share (%)	16.2	19.8	18.9	17.2	16.5	14.8	12.8	11.2	9.3	7.9
Holds gov. bond (%)	1.0	2.7	5.6	9.2	12.8	19.2	25.4	32.7	38.9	41.9
Gov. bonds (%)	0.6	1.2	3.6	5.4	7.6	10.9	14.5	18.3	21.9	23.1
Equity (%)	92.1	92.1	88.9	87.1	84.2	80.2	76.2	71.6	66.2	62.6

Notes: Deciles of V_{it} formed within observation date, then pooled. "Single-asset" is the share of investor-years holding exactly one position. "Domestic", "Gov. bonds" and "Equity" are mean investor-level value shares, and "Holds ..." rows are participation rates. "Fund/ETF value share" is the mean share of own portfolio value in investment funds and exchange-traded funds, the unconditional investor-level portfolio-value share f_{it} in Equation (6), including zero allocations among nonparticipants.

Table 9 Cross-sectional estimates with respect to portfolio value

Outcome	b(ln value), year FE	SE	R2	b(ln value), +demographics	SE	R2	N
N of assets	1.593***	(0.024)	0.178	1.497***	(0.024)	0.200	96,220
$\ln(N_{\text{eff}})$	0.135***	(0.002)	0.139	0.133***	(0.002)	0.179	96,220
HHI	-0.053***	(0.001)	0.132	-0.053***	(0.001)	0.171	96,220
Domestic share	-0.012***	(0.001)	0.018	-0.008***	(0.001)	0.029	96,220
Holds fund/ETF	0.031***	(0.001)	0.037	0.036***	(0.001)	0.074	96,220
Holds foreign asset	0.052***	(0.001)	0.069	0.045***	(0.001)	0.090	96,220

Notes: Equation (2) and its analogues on the pooled panel of 96,220 investor-years. All specifications include observation date fixed effects; the second set adds indicators for sex, age group, education, income band and reported experience, including the non-disclosure categories. Rows three to six report β_s of (7) and β_p of (6). Standard errors clustered by investor in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Portfolio value is jointly determined with composition, so the coefficients summarise the cross-sectional gradient.

Table 10 Investor-characteristic differences in portfolio breadth: unadjusted and conditional estimates

Dimension	Category	Raw gap	SE	Gap | ln value	SE	Share explained (%)	N
Sex	F	-1.897***	(0.094)	-1.417***	(0.087)	25.3	95,056
Age	-18	-0.521	(2.514)	-1.985	(2.462)	-281.2	95,052
Age	18-24	-1.869***	(0.158)	-0.249	(0.154)	86.7	95,052
Age	35-44	1.229***	(0.105)	0.047	(0.094)	96.2	95,052
Age	45-54	1.022***	(0.130)	-0.572***	(0.116)	156	95,052
Age	55-64	0.813***	(0.221)	-0.989***	(0.197)	221.6	95,052
Age	65+	-0.456*	(0.264)	-2.529***	(0.249)	-454.2	95,052
Education	$<$ 10 grades	-1.261	(0.859)	-0.291	(0.823)	76.9	64,216
Education	High school	-0.958***	(0.254)	0.079	(0.237)	108.3	64,216
Education	Post-secondary non-tertiary education	-1.739***	(0.454)	-1.452***	(0.439)	16.5	64,216
Education	Postgraduate	0.088	(0.140)	-0.063	(0.129)	171.6	64,216
Income	6,000-12,000 EUR	0.726***	(0.222)	0.279	(0.206)	61.6	64,198
Income	12,000-24,000 EUR	1.341***	(0.195)	0.381**	(0.183)	71.6	64,198
Income	24,000-48,000 EUR	2.000***	(0.197)	0.390**	(0.183)	80.5	64,198
Income	48,000-120,000 EUR	2.838***	(0.243)	0.349	(0.223)	87.7	64,198
Income	more than 120,000	2.042***	(0.450)	-2.117***	(0.413)	203.7	64,198
Experience	6-12 months	-0.612***	(0.196)	-0.183	(0.179)	70	49,849
Experience	1-3 years	-0.245	(0.247)	-0.631***	(0.225)	-157.2	49,849
Experience	more than 3 years	-0.485**	(0.232)	-2.141***	(0.218)	-341.8	49,849

Notes: Equations (3) and (4); the dependent variable is N_{it}. Each dimension is estimated on the subsample reporting that characteristic, with base categories in the column header. Both specifications include date fixed effects and the second adds $\ln V_{it}$. Standard errors clustered by investor. "Share explained" is the coefficient attenuation, calculated as $100(1 - \tilde{\gamma}_k/\gamma_k)$. It describes changes in estimated coefficients when log brokerage portfolio value is included. It is not a causal mediation measure and is unstable when the unadjusted coefficient approaches zero. Negative values indicate an increase in coefficient magnitude when the coefficients retain the same sign; values exceeding 100% indicate a sign reversal. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

tributions of these mechanisms or establish a causal relationship between portfolio value and the number of distinct financial instruments held.

5.3 Delegation

Table 12 documents rapid growth in the use of investment funds and exchange-traded funds. The share of investors holding at least one such financial instrument rises from 29.1% in 2020 to 51.4% in 2023. Over the same period, the value-weighted share of these instruments in aggregate holdings rises from 4.9% to 7.8%. Among investors who hold at least one fund or exchange-traded fund, these instruments account for 31.2% of portfolio value in 2023, compared with 24.8% in 2020. These measures capture three distinct dimensions: participation, the share of aggregate client portfolio value invested in funds and ETFs, and the average investor-level allocation among fund/ETF holders. They should not be interpreted as interchangeable measures of delegated diversification.

The cross-section separates two distinct patterns. The observed share of investors of holding at least one fund or exchange-traded fund rises with portfolio value from 26.0% in the first value decile to 53.8% in the seventh decile, and then falls to 49.9% in the tenth. In contrast, the share of portfolio value held in these instruments increases from 16.2% in the first decile to 19.8% in the second, before declining to 7.9% in the tenth decile (Table 8). Thus, fund/ETF participation and the unconditional portfolio-value share exhibit different relationships with portfolio value. Neither relationship is strictly monotonic over the full decile distribution.

These two margins correspond to the two equations in (6). Joint estimation yields an unconditional portfolio-allocation slope of $\hat{\beta}_f = -0.0121$ (standard error 0.0006) and a participation slope of $\hat{\beta}_p = 0.0306$ (standard error 0.0010), with correlation $\hat{\rho} = 0.62$ between the estimated slope coefficients (Table 11). Both are precisely estimated in the predicted directions. The negative allocation coefficient and positive participation coefficient provide statistical support for the joint directional pattern specified in H2. However, they do not establish that the observed portfolio choices arise from fixed information costs or delegated diversification rather than other differences across investors.

In economic terms, a doubling of portfolio value is associated with a 0.8 percentage point lower fund/ETF portfolio-value share relative to a mean of 12.4%, and a 2.1 percentage point higher probability of holding any fund or exchange-traded fund relative to a mean of 47.6%. The estimated relationships concern relative allocation and participation, respectively. A lower fund/ETF share in a larger portfolio does not necessarily imply a lower absolute amount invested in these financial instruments.

The decline in participation in the top two portfolio-value deciles may be related to differences in portfolio preferences, financial circumstances, product selection, or the composition of the largest accounts, but these explanations cannot be distinguished using the current

evidence. Additionally, within-investor estimates of the unconditional fund/ETF allocation margin are economically small. In the sample of investors observed at least twice, the estimated coefficient is −0.008, compared with 0.001 in the balanced panel. The allocation pattern should therefore be read primarily as a cross-sectional regularity.

These results indicate that the within-investor association depends on the estimation sample. The findings also illustrate why fund and ETF ownership should be distinguished from portfolio breadth and from the diversification of underlying investment risk. A fund or ETF is counted as one financial instrument, irrespective of the number and composition of the assets it holds. Consequently, portfolio-breadth measures do not capture the underlying exposures of these financial instruments. Table A3 in the Appendix shows that 23.3% of investor-date observations containing only one distinct financial instrument hold a fund or exchange-traded fund. For these observations, the only financial instrument held is a fund or ETF, although the underlying degree of diversification cannot be established from the brokerage data. The share rises to 82.2% among those holding more than 20 distinct financial instruments. Thus, fund/ETF participation is positively associated with total portfolio breadth. However, because portfolio breadth includes the funds and ETFs themselves, this relationship does not establish that fund ownership complements direct holdings or rules out substitution between pooled investment vehicles and individual securities.

Table 11 Joint tests for H2 and H3

Test	Estimate	SE	t	Decision at 5%
H2, unconditional fund/ETF allocation β_f	-0.0121	0.0006	-21.1600	reject $\beta_f = 0$
H2, fund/ETF participation β_p	0.0306	0.0010	31	reject $\beta_p = 0$
H2 joint, Wald $\chi^2(2)$	3,645			reject
H3, within-investor coefficients gap $\eta_N - \lvert\eta_s\rvert$	0.5334	0.0271	19.7000	reject equality

Notes: The first two rows report the estimated slopes of the unconditional fund/ETF portfolio-value share and the indicator for holding at least one fund or ETF, respectively. Both equations include all eligible investor-date observations, including investors with no fund or ETF holdings. The Wald statistic tests the joint null that both coefficients equal zero. The intersection–union procedure tests whether the allocation coefficient is negative and the participation coefficient is positive. These tests evaluate the specified directional relationships but do not identify the economic mechanisms generating them. Standard errors are clustered by investor. The last row refers to the within-investor estimates of (5) and (7) and tests $H_0 : \eta_N = |\eta_s|$ against $H_1 : \eta_N > |\eta_s|$ by the delta method, where η denotes a response standardized by the within-investor standard deviation of the outcome.

Taken together, these results indicate that fund and ETF participation is positively associated with brokerage portfolio value, while the unconditional share of portfolio value allocated to these financial instruments exhibits a negative average relationship with portfolio size. The distinction between participation and allocation is consistent with heterogeneous portfolio-choice mechanisms, including differences in the costs and benefits of delegated investment. However, the evidence does not establish that fund ownership reflects a common diversification motive or that the observed relationships are driven by fixed information costs. In particular, the inclusion of a prominent domestic closed-end fund alongside ETFs,

Table 12 Fund and ETF ownership, portfolio allocation, and government bonds

	2020	2021	2022	2023
Investors holding at least one fund or ETF (%)	29.05	44.36	50.73	51.38
Fund/ETF share of aggregate portfolio value (%)	4.93	6.38	7.89	7.81
Mean investor-level fund/ETF allocation among holders (%)	24.83	29.92	30.49	31.24
Holds government bond (% investors)	16.47	12.64	19.73	21.80
Investors holding government bonds (%)	27.96	19.13	24.17	28.07
Mean government-bond allocation among holders (%)	60.83	52.17	55.01	57.87
Distinct government-bond lines held by sampled investors	6	17	27	31

Notes: The fund/ETF category combines financial instruments classified as ETFs and investment funds, including Fondul Proprietatea. The aggregate allocation measure is weighted by portfolio value, whereas the allocation among holders is the equally weighted mean of investor-level fund/ETF portfolio shares conditional on participation. These measures do not incorporate the underlying holdings of funds and ETFs and therefore cannot directly quantify the economic diversification achieved through these financial instruments. "Distinct government bonds" counts Ministry of Finance lines with at least one holder.

and the absence of information on underlying fund holdings, limit our ability to interpret these patterns as evidence of delegated diversification.

5.4 Domestic portfolio allocation and home bias

Table 13 reports the distribution of brokerage portfolio value across financial instruments classified by issuer country. The value-weighted share invested in Romanian instruments is 85.0%, 85.9%, 87.9%, and 87.5% across the four observation dates. The equal-weighted mean across investors rises from 87.7% to 93.0%. The median investor holds a portfolio consisting entirely of financial instruments classified as Romanian in every year. These measures characterize the country classification of financial instruments held through the observed broker. They do not necessarily capture the geographic distribution of underlying assets or economic risk exposures, particularly for investment funds and ETFs.

Most investors hold no foreign financial instrument. Between 69% and 74% of investors hold only financial instruments classified as Romanian, and the share holding any foreign asset falls from 30.1% in 2020 to 25.8% in 2023. Appendix Table A9 shows that foreign holdings are concentrated in financial instruments associated with a small number of issuer countries. Outside Romania, only the United States (5.9–8.6% of value) and Germany (1.9–2.8%) account for material shares, and together they make up roughly three-quarters of the portfolio value allocated to foreign-classified financial instruments at each date.

The domestic share varies little with the number of instruments held. Investors holding one instrument are 91.9% domestic, and investors holding 11 to 20 instruments are 91.7% domestic. The domestic share falls to 86.7% only among investors holding more than 20 instruments (Appendix Table A3). The portfolio-value gradient across investors is also negative, although its economic magnitude depends on the range of portfolio values considered. The estimated slope is $\hat{\beta}_s = -0.012$ (standard error 0.001), which corresponds to a 0.8

percentage point decline in the domestic share per doubling of portfolio value. Across value deciles, the domestic share falls from 94.8% in the first decile to 85.6% in the tenth (Table 8, Appendix Table A4). Thus, even the highest portfolio-value decile of this broker's clientele, which has access to international markets through the platform, remains about 86% domestic.

Hypothesis 3 concerns within-investor change. The within-investor coefficient on $\ln V_{it}$ is $\hat{\beta}_s^W = -0.0054$ for the domestic share. This corresponds to an estimated decrease of approximately 0.37 percentage points in the domestic portfolio-value share for a doubling of brokerage portfolio value. The corresponding within-investor coefficient for portfolio breadth is $\hat{\beta}^W = 1.780$. Because the two dependent variables are measured in different units, the raw coefficients are not directly comparable.

To compare magnitudes, each response is standardized by the within-investor standard deviation of the outcome. The within-investor standard deviations are 2.917 instruments for N_{it} and 0.070 for the domestic share. The standardized responses are $\hat{\eta}_N = 0.610$ and $\hat{\eta}_s = -0.077$, so the absolute standardized domestic-share coefficient is approximately 12.6% of the standardized portfolio-breadth coefficient. This comparison describes the relative magnitudes of two standardized regression associations. It does not mean that domestic allocations change by 12.6% of the number of financial instruments added, nor does it establish that the two outcomes respond to a common exogenous change in wealth.

Portfolio value, breadth, and domestic allocation are jointly determined by the composition and valuation of brokerage holdings. Acquiring a previously unheld financial instrument increases portfolio breadth, but it does not necessarily increase portfolio value if the purchase is financed by reducing an existing position. Similarly, the domestic portfolio-value share is a ratio involving domestic holdings and total portfolio value, so changes in relative asset prices and portfolio composition can affect both the dependent variable and the regressor. The standardized-coefficient comparison therefore cannot be interpreted as isolating behavioral responses to portfolio growth. We therefore treat the equivalence test in Table 14 as the primary assessment of whether the within-investor association between portfolio value and domestic allocation is economically small under a prespecified equivalence margin.

The equivalence test points in the same direction. Against a margin of one percentage point of domestic share per log point of portfolio value, the null $H_0 : |\beta_s^W| \geq 0.01$ is rejected at the 5% level in both the balanced panel and the larger sample. At the estimated rate, a doubling of portfolio value is associated with a decline of approximately 0.35 percentage points in the domestic share in the larger sample.

Table 15 reports corresponding levels for the balanced panel of 8,718 investors present at all four observation dates. Mean breadth rises from 7.32 to 10.95 and mean N_{eff} rises from 3.90 to 5.06. Fund participation rises from 30.2% to 43.5%, and the median portfolio value rises from RON 28,744 to RON 66,345. Over the same period, the domestic share moves only from 88.1% to 88.5%. hus, investors continuously observed in the brokerage sample exhibit

substantial increases in average portfolio breadth, the mean effective number of holdings, fund/ETF participation, and median nominal portfolio value, while their mean domestic portfolio-value share changes comparatively little. This descriptive evidence is consistent with the small within-investor domestic-share coefficients, although it does not establish that individual investors maintain unchanged allocations or identify the mechanisms underlying portfolio adjustment.

Table 13 Domestic portfolio allocation and foreign-instrument participation

	2020	2021	2022	2023
Domestic share, value-weighted (%)	85.03	85.91	87.89	87.50
Domestic share, mean investor (%)	87.69	89.91	91.06	92.98
Domestic share, median investor (%)	100	100	100	100
Investors 100% domestic (%)	69.94	69.88	69.42	74.23
Investors holding any foreign asset (%)	30.12	30.15	30.62	25.81
D2 sensitivity: max change (pp)	0.13	0.06	0.06	0.07

Notes: Value-weighted figures aggregate over all positions; mean and median investor figures are equal-weighted across investors. The last row reports the maximum possible change in the value-weighted domestic share if every symbol whose country assignment relies on the coarse fallback field were reclassified as foreign, bounding the classification uncertainty discussed in Section 4.

Table 14 Equivalence tests of the within-investor domestic portfolio-share gradient

Sample	$\hat{\beta}_s^W$	SE	t (lower)	t (upper)	Equivalence at 5%
Investors seen ≥ 2 times	-0.0050	0.0010	5	-15	yes
Balanced panel (all four)	-0.0030	0.0020	3.5000	-6.5000	yes

Notes: Two one-sided tests (Schuirmann, 1987) of $H_0 : |\beta_s^W| \geq \delta$ against $H_1 : |\beta_s^W| < \delta$ with $\delta = 0.01$, one percentage point of domestic share per log point of portfolio value, fixed before estimation. The lower statistic tests $\beta_s^W \leq -\delta$ and the upper tests $\beta_s^W \geq \delta$; equivalence requires the first to exceed 1.645 and the second to fall below -1.645.

Taken together, these results document a persistent predominance of Romanian-classified financial instruments in the observed brokerage portfolios. Although investors with larger portfolios allocate a somewhat smaller share to domestic instruments, the corresponding within-investor association is economically small, and the mean domestic share remains relatively stable among investors observed at all four dates. These findings are consistent with the potential relevance of familiarity, information costs, and other investor-specific portfolio-choice frictions. However, the data do not distinguish among these mechanisms, and access to international markets through the brokerage platform does not imply that all investors face identical effective investment opportunities. Moreover, issuer-country classifications do not necessarily capture the underlying geographic exposures of investment funds and ETFs. The results should therefore be interpreted as evidence on domestic portfolio allocation rather than as a comprehensive measure of international risk diversification.

Table 15 Portfolio Characteristics of investors; balanced panel

	2020	2021	2022	2023
Investors	8,718	8,718	8,718	8,718
Mean portfolio breadth N	7.32	9.90	10.24	10.95
Median N	5	7	7	8
Mean effective number of holdings N_{eff}	3.90	4.78	4.86	5.06
Mean HHI	0.48	0.43	0.42	0.41
Domestic share (%)	88.14	88.07	87.97	88.52
Gov.-bond share (%)	10.14	7.47	9.60	11.00
Holds fund/ETF (%)	30.21	36.69	39.97	43.47
Median value (RON)	28,744.05	52,154.91	53,178.83	66,345.37

Notes: 8,718 investors observed at all four observation dates. Domestic share and government-bond share are mean investor-level value shares in percent; "holds fund/ETF" is the participation rate.

5.5 Observed investment opportunities and the aggregate client portfolio

Table 17 counts the distinct trading symbols associated with financial instruments held by at least one client during our sample. Clients held 674 Romanian financial instruments identified by distinct trading symbols, including 380 equities, 43 government bond lines, 84 corporate bonds, 84 certificates, and 4 exchange-traded funds. They also held 3,534 foreign instruments, including 1,838 equities and 335 exchange-traded funds. Foreign instruments therefore outnumber domestic ones by about five to one among the distinct trading symbols observed in client holdings across the four observation dates. The median investor holds five distinct financial instruments at a given observation date, substantially fewer than the 4,208 distinct trading symbols held across the entire client base during the sample period.

The platform-wide count demonstrates access to a broad range of domestic and foreign financial instruments within the observed client population. However, it does not imply that all 4,208 instruments were simultaneously available or equally accessible to every investor. Two considerations qualify this evidence. First, the broker markets international access, so its clientele may be selected on wanting it. As a result, M_t, the observed range of foreign financial instruments may not characterize the range available through other Romanian intermediaries or the investment opportunities effectively facing Romanian households generally. Second, the count of distinct trading symbols aggregates across 43,360 investors and four observation dates, and therefore describes financial instruments held within the client population rather than the precise opportunity set available to any individual investor at a given date. These qualifications also apply to the domestic trading-symbol counts: the number of domestic instruments observed in brokerage holdings is not necessarily the number effectively accessible to each investor. The findings therefore establish the coexistence of a relatively broad observed range of financial instruments and narrow individual portfolios, without ruling out investor-specific investment constraints.

The relationship between the date of first appearance in the brokerage sample and foreign-instrument participation provides additional descriptive evidence on differences across investor cohorts. However, these comparisons cannot establish whether selection into this broker has strengthened or weakened over time, because cohort differences may also reflect portfolio value, investor experience, account tenure, or other unobserved characteristics.

Measured in 2023, foreign participation is 35.3% among investors who first appear in 2020 and 13.7% among those who first appear in 2023. Part of this gap is associated with differences in brokerage portfolio value. The 2023 cohort has a median portfolio of RON 5,918, compared with RON 64,009 for the 2020 cohort, and foreign participation rises steeply with portfolio value. Controlling for $\ln V_{it}$ reduces the gap but does not eliminate it. The 2023 cohort remains 13.0 percentage points below the 2020 cohort (standard error 0.6). The persistence of a conditional difference indicates that portfolio value alone does not account for variation in foreign participation across the first-observation cohorts. However, this comparison does not distinguish differences in investor preferences from differences in account tenure, experience, access-related costs, or other unobserved characteristics.

The client base may still be more internationally oriented than Romanian households as a whole, and nothing internal to these data can settle that.

Table 16 Foreign-instrument participation in August 2023 by first-observation cohort

	2020	2021	2022	2023
Investors in August 2023	8,958	8,162	9,157	14,957
Holds at least one foreign instrument (%)	35.3	36.4	27.0	13.7
Median brokerage portfolio value (RON)	64,008.8	28,581.0	15,551.2	5,917.7

Notes: Investors observed at the 2023 observation date, grouped by the first date at which they appear. Foreign participation is the share holding at least one non-Romanian instrument. Cohort gaps relative to the 2020 cohort, raw and conditional on $\ln V_{it}$, are in Appendix Table A6.

Table 18 provides a complementary comparison based on the portfolio obtained by aggregating the holdings of all observed clients at each observation date. Aggregating all holdings into a single value-weighted portfolio, the effective number of holdings is 17.4 in 2020 and 44.0 in 2023. The corresponding mean individual effective numbers of holdnings are 3.76 and 4.11. In 2023 the client base collectively held 3,160 distinct financial instruments. Of these, 15 instruments account for half of aggregate value and 148 account for ninety percent. The aggregate effective number of holdings is a concentration measure calculated from the combined client holdings, whereas the mean individual effective number is an average of separate investor-level concentration measures. These statistics describe different objects and should not be interpreted as alternative measures of the diversification available to an individual investor.

The aggregate portfolio becomes substantially broader in concentration-equivalent terms, while average individual portfolio concentration changes relatively little. The aggregate

effective number of holdings can also be compared descriptively with the distribution of individual effective numbers. In 2023, the aggregate effective number of 44.0 lies at the 99.97th percentile of the individual N_{eff} distribution. Only 13 investors out of 41,234 reach it. The 99.9th percentile is 31.9 and the maximum is 74.6. These comparisons show that the aggregate client portfolio has a higher effective number of holdings than nearly all individual portfolios. However, the position of the aggregate statistic within the individual distribution does not establish that the aggregate portfolio was practically attainable by each investor.

The aggregate portfolio is an accounting construct rather than an investment portfolio held by any individual client. Replicating its exact financial-instrument weights would require holding the financial instruments included in the aggregate portfolio and would be subject to transaction costs, minimum investment amounts, and other practical constraints. However, achieving an effective number of holdings of approximately 44 does not require replicating the aggregate portfolio or holding thousands of distinct financial instruments. The aggregate comparison therefore documents the concentration structure of combined client holdings without establishing an individual attainability bound.

The observed range of domestic and foreign financial instruments indicates that the broker's clients collectively invest in a substantially broader set of financial instruments than is represented in individual portfolios. This finding weakens explanations that attribute narrow portfolios solely to the absence of financial instruments within the observed brokerage environment. However, it does not rule out supply-side constraints affecting individual investment decisions, including minimum investment amounts, transaction costs, limited liquidity, and differences in effective access to financial instruments.

Demand-side mechanisms, including information costs, attention, familiarity, beliefs, and preferences for particular return distributions, may also contribute to portfolio concentration (Mitton and Vorkink, 2007; Barberis and Huang, 2008; Kumar, 2009). These mechanisms are not mutually exclusive, and the present comparisons do not identify their relative importance.

Supply conditions still shape where investors place their holdings, and the aggregate client portfolio remains concentrated. The ten largest holdings account for 41.6% of aggregate value in 2023, down from 54.1% in 2020. Several of the largest financial instruments are issued by the Romanian government or by companies in which the state retains substantial ownership interests (Table 19). The state is the issuer of the largest bond lines and the controlling shareholder of several large listed firms, including Hidroelectrica, Nuclearelectrica, Romgaz, and Transgaz. Fondul Proprietatea, a fund established in connection with Romania's post-communist property restitution process, also occupies a prominent position in the aggregate client portfolio.

Sector concentration is also high. Utilities, manufacturing (including the integrated oil and gas producer), finance and insurance, transportation, and mining together account for 72.0% of the value of domestic equity held in 2023 (Table 20). Utilities alone rise from

12.4% to 19.8% between 2020 and August 2023, a period that includes the Hidroelectrica listing.

Excluding the largest accounts has little effect on investor-weighted portfolio breadth and concentration, but the composition of aggregate holdings is more sensitive, particularly for government bonds (Appendix Table A7).

These figures indicate that the domestic equity holdings observed in the brokerage sample are concentrated in a relatively small number of reported sectors. However, the sector classifications do not directly measure correlations among securities or the number of independent sources of portfolio risk. Moreover, the classification scheme assigns economically related energy companies to several different sectors, potentially obscuring their common exposures.

An important change in the observed domestic investment opportunity set is the expansion of exchange-traded government bonds directed toward retail investors. Government bonds are the second-largest financial-instrument category by aggregate portfolio value throughout. Their value-weighted share is 28.0% in 2020, 19.1% in 2021, and 24.2% and 28.1% in 2022 and 2023 (Table 12; see also Appendix Table A10, Table A8, and Figure A1). The number of distinct Ministry of Finance bond lines held by clients rises from 6 to 17 to 27 to 31 across the observation dates. This expansion documents an increase in the variety of government-bond lines observed in client holdings. It does not, by itself, establish how much of the change reflects new issuance, differences in investor demand, or the availability and maturity of particular bond lines.

By 2023, one government bond line is the single largest holding in the aggregate client portfolio, and three of the ten largest holdings are government bonds. Participation rises from 16.5% of investors in 2020 to 21.8% in 2023, and from 1.0% to 41.9% across the first and tenth portfolio-value deciles. Across those same portfolio-value deciles, the mean investor-level share allocated to government bonds rises from 0.6% to 23.1%. Among investors who hold government bonds, these bonds account for 57.9% of the portfolio in 2023 and 60.8% in 2020. Holding a government bond is therefore typically associated with a substantial portfolio allocation rather than a negligible position.

The recovery in government-bond holdings in 2022–2023 coincides with both a period of elevated inflation and the expansion of retail-oriented government-bond issuance. With four observation dates and no transaction data, this co-movement can be described but cannot be decomposed into issuance, revaluation, and reallocation. Consequently, the observed increase in government-bond holdings cannot be attributed to a particular policy intervention or to investor responses to inflation without additional evidence.

Taken together, these findings reveal a substantial difference between the range of financial instruments observed across the brokerage client base and the much narrower portfolios held by individual investors. The presence of numerous domestic and foreign instruments indicates that limited platform-wide availability alone cannot fully account for the observed

Table 17 Trading symbols in the brokerage records and client holdings, 2020–23

Instrument type	RO: in file	RO: held	Foreign: in file	Foreign: held
Bonds	84	84	9	9
Certificate	84	84	209	209
ETC	0	0	2	2
ETF	4	4	336	335
Government bonds	43	43	0	0
Investment fund	1	1	0	0
N/A	76	67	1,150	1,140
Rights	42	11	12	1
Shares	380	380	1,844	1,838
TOTAL	714	674	3,562	3,534

Notes: "In file" counts distinct trading symbols recorded in the broker's financial-instrument file. "Held" counts distinct trading symbols appearing in at least one nonempty client holding at one or more of the four observation dates. The counts are pooled across dates and do not represent the number of financial instruments simultaneously available to individual investors. The category "N/A" denotes missing financial-instrument type classification, not necessarily missing issuer-country information.

Table 18 Aggregate client holdings and individual portfolios

	2020	2021	2022	2023
Distinct securities held	2,093	2,694	2,892	3,160
Top-5 share of value (%)	39.62	30.43	29.41	25.32
Top-10 share of value (%)	54.09	43.77	42.33	41.64
Top-20 share of value (%)	67.47	56.15	55.79	56.24
Securities for 50% of value	9	15	15	15
Securities for 90% of value	100	146	151	148
N_{eff} of aggregate portfolio	17.40	34.58	38.04	43.98
Mean investor N_{eff}	3.76	4.50	4.46	4.11

Notes: The aggregate portfolio is the value-weighted sum of all client holdings at the observation date, and its N_{eff} is computed over the security weights in that aggregate. "Mean investor N_{eff}" is the simple average of individual effective numbers, repeated from Table 3. The ratio of the two rows is descriptive. H4 is assessed by locating the aggregate within the empirical distribution of individual effective numbers, reported in the text.

Table 19 Ten largest financial instruments by aggregate client holdings value, August 2023

Symbol	Issuer	Instrument	Country	Share (%)
R2408A	Ministry of Finance	Government bonds	Romania	6.94
TLV	Banca Transilvania	Shares	Romania	6.07
H2O	Hidroelectrica	Shares	Romania	4.12
SNP	OMV Petrom	Shares	Romania	4.10
FP	Fondul Proprietatea	Investment fund	Romania	4.09
SNN	Nuclearelectrica	Shares	Romania	4.03
R2808AE	Ministry of Finance	Government bonds	Romania	3.52
SNG	Romgaz	Shares	Romania	3.35
R2804AE	Ministry of Finance	Government bonds	Romania	3.33
TGN	Transgaz	Shares	Romania	2.08

Notes: Share of the aggregate value of all client holdings at the 22 August 2023 observation date. Symbols beginning with R are Ministry of Finance bond lines listed on the Bucharest Stock Exchange. Issuer names are shortened for display.

Table 20 Sector composition of domestic equity holdings (value shares, %)

Sector (LSEG block of the enriched symbols sheet)	2020	2021	2022	2023
Utilities	12.4	12.6	12.5	19.8
Manufacturing	15.6	19.0	19.8	18.5
Finance and Insurance	21.9	18.2	19.5	18.4
Transportation and Warehousing	10.9	6.7	8.6	8.1
Mining, Quarrying, and Oil and Gas Extraction	8.6	6.6	7.5	7.2
Professional, Scientific, and Technical Services	6.3	7.6	7.3	6.7
Information	0.2	2.0	4.1	4.5
Wholesale Trade	10.4	8.2	5.7	4.1
All other sectors	13.7	19.2	15.0	12.7
TOTAL	100.0	100.0	100.0	100.0

Notes: Sector shares describe the allocation of observed domestic equity holdings across the reported classification categories. They do not measure the number of independent risk exposures or the covariance structure of the underlying securities. Value-weighted shares of the domestic equity block only, excluding government bonds, corporate bonds, funds and foreign assets. Sector labels come from the commercial-data enrichment block of the instrument file and mix two classification vocabularies; the assignment for the largest issuers was inspected manually, while the long tail of small AeRO issuers should be treated as indicative. The integrated oil and gas producer is classified under Manufacturing and the gas transmission operator under Transportation, so the energy complex is split across four of the five leading rows.

portfolio concentration. However, the holdings data do not establish that all investors face identical effective investment opportunities or that supply-side constraints are unimportant. Market liquidity, transaction costs, minimum investment requirements, and investor-specific information and preferences may jointly shape portfolio choices. The expansion of retail-oriented government bonds further illustrates how changes in the domestic investment opportunity set can influence the composition of brokerage holdings. The evidence therefore supports a framework in which institutional market conditions and investor-level portfolio-choice mechanisms operate jointly, rather than as mutually exclusive explanations of narrow portfolios.

5.6 Entry, exit, non-disclosure and the censored final date

Between the 2022 and 2023 observation dates, the mean number of instruments held falls by 0.58 and the equally weighted mean of investor-level domestic portfolio shares rises by 1.9 percentage points. The decomposition identity in Equation (8) shows that both attributable to changes in the composition of the observed investor population, particularly the appearance of investors with narrower and more domestically concentrated portfolios.

Table 21 applies the decomposition. Of the 41,234 investors observed in 2023, 26,004 were present in 2022, 15,230 observed in 2023 but absent in 2022, and 939 investors present in 2022 are no longer observed in August 2023. Continuing investors increased the number of instruments held by 1.185 on average. Entrants hold 4.48 instruments on average, compared with 9.70 for continuing investors, so the entry-composition component contributes - 1.929 instruments to the change in the cross-sectional mean. Exit contributes +0.162 because

investors absent from the later snapshot held fewer instruments at the earlier date than those observed at both dates. The three components sum to the observed change of −0.582.

Raw entrant counts are not directly comparable across observation intervals of different lengths because the observation windows are not equally long. The first two windows cover 365 days and the third covers 234 days. Thus, the 15,230 entrants observed between 2022 and 2023 first appear relative to the preceding snapshot over roughly two-thirds of a year, while the 9,749 entrants between 2021 and 2022 are observed over a full year. Table 22 adjusts for this difference. As a share of the investor base at the start of the window, entry is 90.7% in 2020–2021, 53.5% in 2021–2022, and 56.5% in 2022–2023. Under a proportional adjustment for interval length, the annualized entry rate for the final window is 88.2%, close to the 2020–2021 rate. In annualized count terms, the final window corresponds to 23,756 entrants per year, compared with 9,749 in the previous window. By contrast, the similarly annualized exit rate varies relatively little, between 5.4% and 5.6% of the base in all three windows. These annualized figures rescale observed changes according to the duration of each interval. They should not be interpreted as observed calendar-year entry and exit rates or as estimates of a constant underlying arrival process.

The pattern is not confined to 2023. In every transition, continuing investors increase the number of distinct financial instruments held (+2.43, +1.00, and +1.18), while entry composition lowers the cross-sectional mean (−1.19, −1.29, and −1.93). The total change can therefore move in the opposite direction from the average change in portfolio breadth among investors observed at both dates. Figure 4 plots the components. In particular, the 2021–2022 and 2022–2023 transitions exhibit declines in overall mean breadth despite positive average changes among continuing investors.

The same arithmetic applies to the equally weighted mean investor-level domestic portfolio-value share. Of its 1.92 percentage point rise between December 2022 and August 2023, 1.82 points come from the entry composition component. The average domestic portfolio-value share is approximately 96% among entrants, compared with 91.2% for continuing investors. Only +0.12 points come from within-investor change (Appendix Table A2). Appendix Table A1 reports the corresponding decomposition for the effective number of holdings. The latter decomposition provides additional evidence that changes in the observed investor population affect measures of portfolio concentration as well as portfolio breadth.

Appendix Table A17 reports the value-weighted counterpart of the composition effect. By value, investors who first appear in 2020 still hold 61.5% of the total in 2023, although they are only 21.7% of investors. By contrast, 2023 entrants are 36.3% of investors but hold only 9.6% of value. These differences indicate that first-observation cohorts have markedly different weights in investor-weighted and portfolio-value-weighted statistics. The earlier cohort contributes disproportionately to aggregate brokerage portfolio value, whereas the most recently observed cohort accounts for a much larger share of investor observations than of value. Consistent with these differences, Table 4 shows that the value-weighted mean

portfolio breadth and effective number of holdings increase across the four dates, while their equally weighted counterparts decline after 2021. This comparison is descriptive and does not itself constitute a formal decomposition of changes in the value-weighted measures.

More generally, during participation waves, cross-sectional averages of retail portfolio breadth and concentration may be strongly influenced by changes in the composition of the investor population. Monitoring frameworks that track average portfolio breadth or concentration without separating entry from within-investor change may therefore misinterpret a decline in an aggregate descriptive statistic as evidence that continuing investors have reduced their portfolio breadth or increased their concentration.

Table 23 suggests that missing information in the investor-characteristics file is systematically associated with observed portfolio characteristics. However, it is not clear whether missing values arise from investors declining to report information, administrative recording practices, or incomplete linkage between data sources.

Investors who reported all characteristics account for 51.8% of investor-years and hold 65.1% of value, with a median portfolio of RON 32,462. The largest incomplete-information group, with education, income, and experience all missing, accounts for 32.0% of investor-years and 7.2% of value, with a median portfolio of RON 6,220. Most of these observations come from 2023: 68% of the group's investor-years are from the final observation date. Simultaneous missingness of all five characteristics is rare (1.2% of investor-years) but accounts for 5.3% of aggregate portfolio value, with a median portfolio of RON 83,455. These differences indicate that the availability of investor characteristics is associated with portfolio value and the timing of observation. They do not, however, establish the reasons for missing information or whether the missingness mechanism is stable over time.

Overall, the observed missing-information patterns are associated with substantial differences in brokerage portfolio value and composition. A large group of observations with missing education, income, and experience information is associated with relatively small portfolios and is disproportionately represented in August 2023. A smaller group with all five characteristics missing is associated with substantially larger portfolios. Consequently, complete-case analysis is not innocuous. It drops about half of investor-years, about one-third of aggregate value, and disproportionately the newest cohort that drives cross-sectional time trends. The implications depend on the variables required by each specification. A regression using only reported sex or age retains considerably more observations than a regression requiring complete information on all five characteristics. Therefore, the available-case samples used in the demographic analysis should be distinguished from the smaller complete-case sample containing all characteristics.

The preceding findings are particularly relevant to the final observation date, which accounts for a substantial share of the pooled sample and differs from the earlier dates in its timing. The August 2023 snapshot is not a year end, it is separated from the preceding date by 234 rather than 365 days, and it is right-censored. As a robustness check, the analysis is

Table 21 Decomposition of the change in the mean portfolio breadth

	2020 → 2021	2021 → 2022	2022 → 2023
Mean, t	6.939	8.404	8.352
Mean, t+1	8.404	8.352	7.770
Total change	1.465	-0.053	-0.582
Within stayers	2.432	0.995	1.185
Entry composition	-1.185	-1.288	-1.929
Exit composition	0.219	0.241	0.162
Stayers	9,282	17,194	26,004
Entrants	8,924	9,749	15,230
Exiters	555	1,012	939
Entrant mean	7.171	6.079	4.477
Stayer mean, t+1	9.590	9.640	9.698
Exiter mean, t	3.279	4.307	3.869
Stayer mean, t	7.158	8.646	8.514

Notes: The table decomposes changes in the equally weighted mean number of distinct financial instruments held between consecutive observation dates, using Equation (8). "Stayers" are investors observed at both dates. "Entrants" are investors present at the later date but absent from the immediately preceding snapshot, and "exiters" are investors present at the earlier date but absent from the later snapshot. These categories describe observed panel participation and do not necessarily identify account openings or closures. The decomposition is an exact accounting identity, subject to numerical rounding. Corresponding decompositions for the effective number of assets and the domestic share are in Appendix Tables A1 and A2.

Table 22 Window lengths, entry and exit rates

	2020 → 2021	2021 → 2022	2022 → 2023
Window (days)	365	365	234
Window (years)	1	1	0.6
Entrants	8,924	9,749	15,230
Entrants per year	8,924	9,749	23,756.2
Entry rate (% of base)	90.7	53.5	56.5
Entry rate, annualised (%)	90.7	53.5	88.2
Exiters	555	1,012	939
Exit rate, annualised (%)	5.6	5.6	5.4

Notes: Entry and exit are defined by presence or absence at consecutive holdings snapshots. Rates use the number of investors observed at the beginning of each interval as the denominator. The annualized figures rescale the observed counts and rates by $365/d$, where d is the interval length in days. They are descriptive adjustments for comparability and do not measure actual calendar-year entry or exit or establish constant underlying arrival and departure rates.

Table 23 Patterns of missing investor characteristics

Missing characteristic(s)	Investor-years	% sample	% value	Median RON	Mean N	Mean N_{eff}	Domestic	Fund/ETF	% in 2023
all reported	49,832	51.8	65.1	32,461.7	9.7	5.0	89.5	51.6	31.5
educ/inc/exp	30,797	32.0	7.2	6,220.1	5.3	3.2	95.2	48.0	68.4
exp	14,331	14.9	22.3	22,219.2	7.8	3.9	89.9	34.7	28.5
sex/age/educ/inc/exp	1,164	1.2	5.3	83,454.9	7.8	4.2	82.3	26.5	29.9
educ/exp	39	0.0	0.0	22,273.2	2.5	1.8	100.0	7.7	25.6
inc/exp	36	0.0	0.0	6,229.8	3.8	2.1	87.1	16.7	22.2
inc	17	0.0	0.0	3,730.8	5.3	4.0	84.9	47.1	35.3
age/educ/inc/exp	4	0.0	0.0	1,809.3	2	1.7	100	0	25

Notes: Each row is a distinct pattern of missing account-opening characteristics and the label lists the blocks that are missing. Only 8 of 32 logically possible patterns occur. "% in 2023" is the share of the pattern's investor-years drawn from the final observation date.

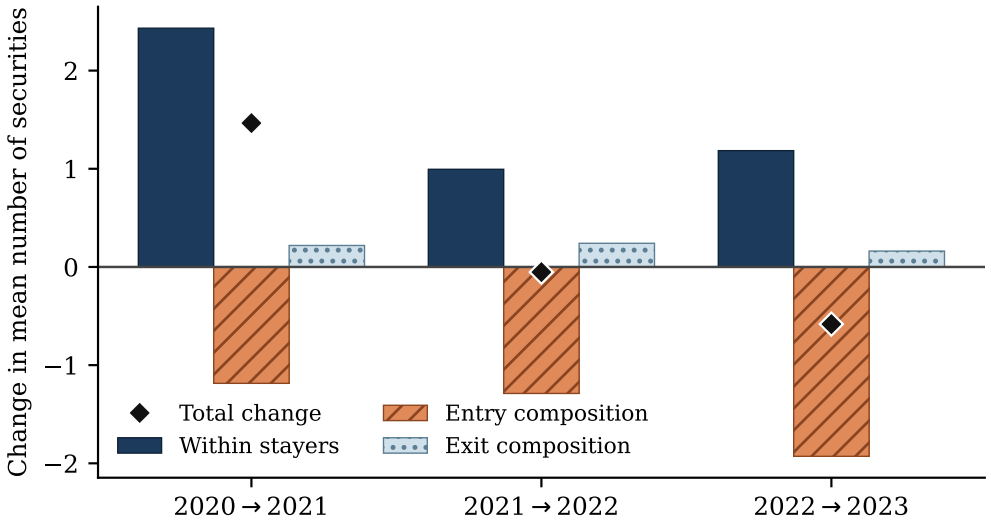


Figure 4 Changes in mean portfolio breadth: continuing investors and sample composition

*Notes:*The figure displays the decomposition of changes in mean portfolio breadth into the mean change among investors observed at both dates and the contributions of investors entering and exiting the observed sample. Components are calculated using Equation (8) for the mean number of positions. Diamonds mark the total observed change.

therefore repeated using only the three year-end dates. This removes 41,234 investor-years (43% of the pooled sample) and, with them, the largest and most recent entrant cohort. It is therefore a sensitivity analysis assessing whether the principal associations depend heavily on including the August 2023 cross-section.

Table 24 reports the value gradients across investors at a given observation date. The gradient in the number of distinct instruments held in (2) is 1.538 on the three year-end dates, compared with 1.593 on all four. The coefficient on $\log N_{\text{eff}}$ is 0.131 versus 0.135. The domestic-share gradient is −0.013 versus −0.012, the share held through funds and exchange-traded funds gradient is −0.009 versus −0.012, and the participation gradient is 0.030 versus 0.031. Each estimate keeps its sign, its statistical significance at the 1% level, and a broadly comparable magnitude, although the fund/ETF allocation coefficient is somewhat smaller in absolute value when August 2023 is excluded.

The within-investor coefficient in Equation (5) is 1.806 on the year-end dates versus 1.780 on all four. Under Hypothesis 3, the comparison of standardized within-investor coefficients gives $\hat{\eta}_N - |\hat{\eta}_s| = 0.545$ with a delta-method standard error of 0.040 and $t = 13.6$, compared with 0.533 and $t = 19.7$ in the full sample. The absolute standardized domestic-share coefficient is approximately 14.6% of the standardized portfolio-breadth coefficient, rather than 12.6%. These comparisons indicate that the main directional portfolio-value associations and the relative standardized coefficient magnitudes remain similar when the August 2023 observation is excluded. However, this exercise does not independently reassess every component of H1b or the descriptive opportunity-set comparison in H4, and it does not identify the mechanisms underlying the estimated relationships.

Table 25 reports levels. The mean number of instruments held is 8.12 on the year-end dates, compared with 7.97 on all four. Mean N_{eff} is 4.35 versus 4.25. Portfolios containing one distinct financial instrument account for 17.6% of investor-years versus 18.1%. The equal-weighted domestic share is 90.1% versus 91.3%, and fund participation is 44.7% versus 47.6%. These differences are consistent with the distinctive composition of the August 2023 cross-section, which contains many newly observed investors with relatively small, narrow, and predominantly domestic brokerage portfolios. Excluding that date changes the pooled descriptive averages but does not eliminate the principal findings of narrow individual portfolios, substantial concentration, and high domestic allocation.

Table 24 Cross-sectional gradients excluding the August 2023 observation date

Outcome	2020-2022	SE	All four dates	SE	N (2020-2022)	N (all)
N of assets	1.538***	(0.028)	1.593***	(0.024)	54,986	96,220
$\ln(N_{\text{eff}})$	0.131***	(0.002)	0.135***	(0.002)	54,986	96,220
Domestic share	-0.013***	(0.001)	-0.012***	(0.001)	54,986	96,220
Fund/ETF value share	-0.009***	(0.001)	-0.012***	(0.001)	54,986	96,220
Holds fund/ETF	0.030***	(0.001)	0.031***	(0.001)	54,986	96,220

Notes: Each row is one estimate of (2) or its analogues, with date fixed effects and standard errors clustered by investor. The first pair of columns uses the three year-end dates, the second all four dates. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 25 Descriptive portfolio statistics: three Year-end observations versus the full sample

	2020-2022	All four dates
Investor-years	54,986	96,220
Mean portfolio breadth N	8.12	7.97
Median N	6	5
Mean effective number of holdings N_{eff}	4.35	4.25
Mean HHI	0.44	0.45
Portfolios containing one financial instrument (%)	17.61	18.12
Domestic, EW (%)	90.08	91.32
Holds fund/ETF (%)	44.74	47.58

Notes: Pooled statistics over the three year-end dates and over all four dates. Differences reflect the composition of the 2023 entrant cohort, which is smaller, narrower and more domestic than the incumbent base.

Taken together, these results demonstrate the importance of changes in the observed investor population for interpreting retail portfolio characteristics. Although mean portfolio breadth and the mean effective number of holdings decline in the full cross-section between December 2022 and August 2023, both measures increase among investors observed at consecutive dates. Similarly, the rise in the equally weighted mean domestic portfolio share is largely associated with the addition of investors holding predominantly Romanian-classified financial instruments. The substantial differences in portfolio value and composition across missing-information patterns further indicate that demographic analyses depend on the

availability of investor characteristics. Excluding the August 2023 observation produces broadly similar portfolio-value associations, suggesting that the principal descriptive relationships are not specific to the final cross-section. These findings highlight the need to distinguish changes within continuing investors from changes in sample composition, while recognizing that the available holdings snapshots do not identify account openings, trading behavior, or the mechanisms generating portfolio adjustments.

6 Discussion

Table 26 summarizes the principal empirical findings associated with H1–H4. The cross-sectional portfolio-value coefficient is positive and precisely estimated (H1a), while several demographic and socioeconomic differences in portfolio breadth become smaller after conditioning on portfolio value (H1b). Fund/ETF participation and the unconditional portfolio-value share exhibit coefficients with opposite signs, consistent with the directional pattern specified in H2. Within investors, the standardized portfolio-value coefficient is substantially smaller for the domestic portfolio share than for portfolio breadth (H3). Finally, the range of financial instruments observed across the brokerage client base is much broader than individual holdings, and the effective number of holdings in the aggregate client portfolio substantially exceeds that of most individual portfolios (H4).

These results establish empirical associations and descriptive comparisons, but they do not uniquely identify the mechanisms motivating the hypotheses. In particular, H4 is not a regression-based parameter restriction and should not be interpreted as a formal rejection of individual investment constraints.

The positive relationship between brokerage portfolio value and portfolio breadth can be considered alongside evidence on household portfolio diversification from Sweden (Calvet et al., 2007), the United States (Goetzmann and Kumar, 2008) and Germany citepvongaudecker2015. However, these studies differ in their portfolio measures, observed assets, and definitions of wealth, preventing direct comparison of coefficient magnitudes.

Home bias is ubiquitous, and no single explanation accounts for its full extent (French and Poterba, 1991; Ardalan, 2019). Gender differences in portfolio behavior are widely documented (Barber and Odean, 2001). The contribution of the present study is to document these dimensions of portfolio choice jointly in a Romanian brokerage setting and to examine how their cross-sectional patterns differ from changes observed within investors and across entry cohorts. The findings also demonstrate that several demographic differences persist after conditioning on brokerage portfolio value, although their magnitudes vary considerably across groups.

Four findings are particularly informative about the coexistence of narrow individual portfolios and a broad observed range of financial instruments. First, the aggregate client

Table 26 Outcome of the hypothesis tests

	Hypothesis	Test statistic	Decision on H_0	Economic magnitude
H1a	The number of distinct instruments held rises with portfolio value	$\hat{\beta} = 1.593$ (0.024) across investors; -0.234 (0.164) within investor, instrumented	Rejected at 1% across investors; not rejected within investor	+1.10 additional instruments per doubling of value across investors
H1b	Demographic differences operate mainly through portfolio value	87–96% of the age gaps and 62–88% of the income gaps absorbed	Rejected at 1% for age, education, income; not rejected for gender	Conditional age and income gaps below 0.4 instruments; women hold 1.42 fewer
H2	Fund and exchange-traded fund shares fall and participation rises with portfolio value	$\chi^2(2) = 3645$; $t_f = -21.2$, $t_p = 31.0$ (intersection–union)	Rejected at 1%	-0.8 pp fund share and +2.1 pp participation per doubling
H3	The domestic share adjusts less than the number of instruments held, within investor	$\hat{\eta}_N - \lvert\hat{\eta}_s\rvert = 0.533$ (0.027), $t = 19.7$	Rejected at 1%	Domestic share moves 12.6% as much as the number of instruments held
H4	The opportunity set does not bind	Not a parameter restriction	Bound at the 99.97th percentile of individual N_{eff} in 2023	13 of 41,234 investors reach it; 3,534 foreign instruments traded

Notes: Standard errors in parentheses. With samples between 35,000 and 96,000 investor-years, nearly every coefficient differs from zero at conventional levels, so the decision and magnitude columns should be read together. H4 carries no sampling distribution; its entry reports where the attainability bound falls in the empirical distribution of individual effective numbers.

portfolio is about ten times larger in effective-number-of-holdings terms than the average client portfolio, although the aggregate statistic does not represent an individually feasible benchmark. Second, financial instruments identified by distinct trading symbols and held by at least one client during the sample period outnumber domestic instruments by roughly five to one. Third, investors holding up to 20 distinct financial instruments exhibit relatively similar mean domestic portfolio shares, while the share is lower among those holding more than 20 instruments. Fourth, the same investors observed at all four observation dates increased their average portfolio breadth, effective number of holdings, and fund/ETF participation, while their mean domestic portfolio share changed relatively little. These findings indicate that the breadth of financial instruments observed across the client base is not reflected proportionally in individual portfolios. They do not establish whether this difference results primarily from investment costs, information constraints, investor preferences, or other characteristics of the market and its participants.

Taken together, these results are compatible with several explanations of retail portfolio concentration, including fixed costs of acquiring and maintaining financial instruments, information costs, familiarity, and heterogeneous investor preferences. The hypotheses were drawn from the recognition and information-acquisition models of Merton (1987) and Van Nieuwerburgh and Veldkamp (2010), the treatment of delegation in Campbell (2006) and von Gaudecker (2015), and the information-immobility and familiarity literature (Coval and Moskowitz, 1999; Huberman, 2001; Van Nieuwerburgh and Veldkamp, 2009). The positive, approximately log-linear relationship between brokerage portfolio value and portfolio breadth is consistent with models in which investors face costs of expanding their holdings. However, the observed functional form is not a unique prediction of fixed information costs and cannot distinguish these costs from other portfolio-choice frictions. As portfolio value rises, the share of wealth held through funds and exchange-traded funds falls on average, while the probability of holding at least one fund or exchange-traded fund rises. This combination suggests that participation and allocation are distinct dimensions of fund ownership. However, it does not identify the underlying investment motives, particularly because the category combines ETFs with a major domestic closed-end investment fund. The persistent predominance of Romanian-classified financial instruments is compatible with familiarity and information-related explanations, but it is also consistent with differences in effective access, transaction costs, investment objectives, and other unobserved investor characteristics. The empirical analysis does not determine the relative importance of these mechanisms.

The gender-related estimates highlight an important difference between the unadjusted and portfolio-value-conditional patterns. Among investors with similar brokerage portfolio value, women hold fewer distinct financial instruments than men. This gap is reduced by approximately one-quarter after conditioning on portfolio value, but remains substantially larger than the corresponding conditional differences for several younger age and middle-

income categories.

However, sizeable conditional differences also remain for older investors, certain education groups, and the highest reported income category. The gender result should therefore not be presented as the only demographic or socioeconomic pattern that persists after accounting for portfolio value.

Market structure may influence both the financial instruments available to investors and the composition of their portfolios. The domestic equity block represented in the brokerage holdings data is concentrated, so holding ten domestic stocks can still imply substantial exposure to common sectoral or issuer-specific risks. The sovereign retail bond program also expanded the range of government-bond lines observed in client holdings and the approximately one-fifth of observed investors holding government bonds in August 2023, represents a substantial share of brokerage portfolio value on average.

These findings illustrate that the number of distinct financial instruments held is only one dimension of portfolio choice. Their economic risk exposures depend on the composition of the instruments, including the correlation of returns across equities, government bonds, and other asset classes. An additional contribution concerns the distinction between changes in individual portfolios and changes in the composition of the observed investor population. The entry-and-exit decomposition shows that average portfolio breadth and the effective number of holdings can decline in the full cross-section even as both increase among continuing investors. Similarly, changes in the equally weighted mean domestic portfolio share are strongly associated with the arrival of investors holding predominantly Romanian-classified financial instruments. These findings demonstrate that cross-sectional changes in retail portfolio characteristics cannot be interpreted as changes in investor behavior without accounting for changes in sample composition.

Three implications follow, framed as conjectures suggested by the descriptive evidence rather than causal or welfare conclusions. First, related to measurement: summary statistics on retail portfolio breadth and concentration can worsen because of changes in the composition of the observed investor population during periods of expanding market participation. Monitoring frameworks at exchanges, regulators, or financial-education programs that track average portfolio breadth or concentration without separating entry from changes among continuing investors may therefore misinterpret broadening participation as deteriorating portfolio characteristics among existing investors. The decomposition in Equation (8) is straightforward to implement using sufficiently consistent investor-level holdings data observed at successive dates.

Second, related to policy design: if narrow portfolios reflect a fixed cost per financial instrument held, encouraging investors to hold additional individual financial instruments may not be the most cost-effective way to improve diversification. A potential alternative is to improve access to appropriately diversified, transparent, and low-cost pooled investment vehicles, while recognizing differences in their underlying exposures and fees. Fund and

exchange-traded fund participation rises from 29% to 51% of investors over three years, demonstrating substantial growth in observed ownership within the brokerage client base. Only approximately a quarter of the smallest value decile holds any fund or exchange-traded fund. The lower participation rate among investors with the smallest brokerage portfolios may reflect financial constraints, investment preferences, product familiarity, fees, or other barriers. The present data do not distinguish among these explanations or establish which policy intervention would be most effective.

Third, the sovereign retail bond program represents a documented expansion in the range of listed government-bond lines held by the observed brokerage clients. Take-up rises from 16% to 22% of clients. With only four observation dates, it is not possible to separate issuance from revaluation from active reallocation, but transaction-level data could. The welfare question of whether households substitute out of deposits, out of equity, or out of neither is important for evaluating the consequences of government-bond programs for household portfolios and capital-market development. Addressing this question would require information on portfolio flows, holdings outside the observed brokerage account, and the investment alternatives available to participating households.

The results characterize the clients of a single full-service domestic intermediary. The sample is conditional on participation through this broker and therefore cannot be assumed to represent all Romanian retail investors or households. The broker's services, fee structure, financial-instrument offerings, and client acquisition practices may influence the characteristics of the observed investor population. However, without comparable data from other intermediaries, neither the direction nor the magnitude of this selection can be established. Consequently, the estimated domestic portfolio shares should not be interpreted as upper or lower bounds for Romanian retail investors generally.

All dynamic statements compare four observation dates. Changes in holdings therefore combine trading, revaluation, corporate actions, and issuance. This matters most for the government-bond share in 2022–2023. More generally, the data do not record transaction histories, net deposits and withdrawals, the timing of portfolio adjustments, or changes in holdings between observation dates. Consequently, the first appearance or disappearance of an investor does not necessarily correspond to brokerage account opening or closure, and changes in portfolio breadth cannot be attributed directly to purchases or sales. The final observation, August 22, 2023, is also separated from the preceding year-end by only 234 days, limiting direct comparison with the earlier annual intervals.

The positive association between brokerage portfolio value and portfolio breadth is observed both across investors and within investors over time. However, portfolio value and portfolio composition are jointly determined, and investor fixed effects do not eliminate time-varying confounding or accounting relationships between the variables. The separate passive-return instrumental-variable specification yields an estimate that is not statistically distinguishable from zero, but its exclusion restriction is not established. Therefore, nei-

ther the fixed-effects estimates nor the passive-return IV results identify a causal effect of portfolio value on breadth. Distinguishing portfolio inertia, investor sorting, and responses to changes in investable resources would require additional information on transactions and cash flows, a longer panel, or a credible source of exogenous variation.

Only one brokerage account is observed rather than a household balance sheet, so statements about portfolio breadth and concentration apply to the observed brokerage holdings. Diversification measures ignore covariance. This can understate the breadth of underlying exposures in portfolios containing diversified funds and ETFs, while overstating the degree of risk diversification obtained from multiple highly correlated financial instruments. Moreover, the effective number of holdings is calculated from the distribution of portfolio value across distinct financial instruments and does not measure the number of independent risk factors or the efficiency of the investor's portfolio. Account-opening characteristics are measured once, are unvalidated for education, income and experience, and are missing in a structured way. Demographic results are therefore restricted to the relevant reporting subsamples and are conditional associations. No causal identification is claimed.

Natural extensions follow from these limitations. Transaction-level data would separate trading from revaluation and permit event studies around the bond issuance calendar and the Hidroelectrica listing. However, such event studies would require an appropriate identification strategy to support causal claims about the effects of these market developments. Linked survey evidence on beliefs, financial literacy, trust and motives would allow the information-cost mechanism behind H1–H3 to be tested more explicitly, although identifying causal mechanisms would still require suitable research designs. Incorporating expected returns and covariances would refine diversification measurement and permit an assessment of mean–variance efficiency. A look-through analysis of investment funds and ETFs would additionally distinguish the number of financial instruments held from the underlying assets and risk exposures represented in investors' portfolios. Extending coverage beyond one intermediary would help settle the representativeness question. Comparable holdings data for Poland, Hungary or Bulgaria would establish whether high observed domestic allocation among brokerage clients is specific to Romania or a broader regional pattern. Further research could also investigate the relationship between market access, transaction costs, and portfolio choices using information on broker-specific fees, minimum investment amounts, liquidity, and the financial instruments available to individual investors. Such evidence would provide a stronger basis for evaluating the relative importance of supply-side and investor-specific constraints.

7 Conclusion

This paper examines the portfolios of 43,360 retail clients of a large Romanian broker, using 96,220 investor-date observations between December 2020 and August 2023. The analysis documents portfolio breadth and concentration, fund and ETF ownership, domestic allocation, and their relationships with brokerage portfolio value and investor characteristics. It also compares individual holdings with the range of financial instruments observed across the client base and separates changes among continuing investors from changes in the composition of the observed investor population.

Individual portfolios contain relatively few distinct financial instruments. The median portfolio holds five instruments, approximately 18% contain only one, and the mean effective number of holdings is 4.25. Brokerage portfolio value is strongly and positively associated with portfolio breadth across and within investors, although the passive-return instrumental-variable analysis does not establish a causal response of breadth to changes in portfolio value. Conditioning on portfolio value substantially reduces several demographic and socioeconomic differences, although meaningful heterogeneity remains. Fund and ETF participation increases with portfolio value on average, while the unconditional share allocated to these instruments exhibits a negative association. Domestic issuer-country concentration is substantial: approximately three-quarters of investors hold only Romanian-classified financial instruments in August 2023, and the mean domestic portfolio share changes relatively little among investors observed at all four dates.

The range of financial instruments observed across the brokerage client base is considerably broader than individual portfolios. In August 2023, the aggregate client portfolio has an effective number of approximately 44 holdings, compared with an individual mean of 4.11. Across the sample period, client holdings include 3,534 distinct foreign financial instruments. These findings demonstrate the coexistence of narrow individual portfolios and a broad observed range of domestic and foreign instruments. However, they do not establish the effective opportunity set facing each investor or distinguish the relative importance of investment costs, information constraints, and heterogeneous preferences.

Changes in the composition of the observed investor population are also important. Between December 2022 and August 2023, mean portfolio breadth declines by 0.58 instruments in the full cross-section, although it increases by 1.19 among continuing investors. The decline is largely attributable to the appearance of investors holding narrower portfolios. Similar composition effects characterize the effective number of holdings and the equally weighted mean domestic portfolio share. Consequently, changes in aggregate retail portfolio statistics need not reflect corresponding changes among continuing investors.

These findings provide a quantitative baseline for understanding retail portfolio choice in Romania's evolving capital market. They highlight the importance of distinguishing portfolio breadth from risk diversification, observed platform-wide holdings from individual in-

vestment opportunities, and changes in investor composition from changes within investors. Although the results are consistent with several mechanisms discussed in household-finance theory, the available data do not identify their relative contributions or establish causal effects. Combining holdings data with transaction histories, investment costs, and information on underlying asset exposures would provide a stronger basis for evaluating the determinants and consequences of retail portfolio concentration.

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Appendix

A Derivation of the decomposition of changes in investor-level means

Let S be the set of investors present at both dates t and $t+1$. Let E be the set present only at $t+1$ (entrants) and X the set present only at t (exiters). These categories refer to investors' presence in successive holdings snapshots and do not necessarily identify actual brokerage account openings or closures. Then $N_t = |S| + |X|$ and $N_{t+1} = |S| + |E|$. For any investor-level statistic b, write $\bar{b}_{G,s}$ for its mean over group G at observation date s. Then

$$\bar{b}_{t+1} = \frac{|S|\bar{b}_{S,t+1} + |E|\bar{b}_{E,t+1}}{N_{t+1}}, \qquad \bar{b}_t = \frac{|S|\bar{b}_{S,t} + |X|\bar{b}_{X,t}}{N_t},$$

and adding and subtracting $\bar{b}_{S,t+1}$ and $\bar{b}_{S,t}$ yields

$$\bar{b}_{t+1} - \bar{b}_t = \underbrace{\left(\bar{b}_{S,t+1} - \bar{b}_{S,t}\right)}_{\text{within}} + \underbrace{\left(\bar{b}_{t+1} - \bar{b}_{S,t+1}\right)}_{\text{entry}} - \underbrace{\left(\bar{b}_t - \bar{b}_{S,t}\right)}_{\text{exit}},$$

where the entry term equals $\frac{|E|}{N_{t+1}}(\bar{b}_{E,t+1} - \bar{b}_{S,t+1})$ and the exit term equals $\frac{|X|}{N_t}(\bar{b}_{X,t} - \bar{b}_{S,t})$. This is equation (8). It is an identity and holds exactly, up to numerical precision, in the reported tables.

The decomposition is of the Foster–Haltiwanger type adapted to a population mean rather than a share-weighted aggregate. The "within" term captures changes among continuously observed investors. The data do not support separating active trading from revaluation, so this decomposition does not attempt to do so. The entry and exit terms are composition effects arising from differences between investors present at only one observation date and those present at both. The decomposition is an exact accounting identity and does not identify the causes of changes in investor participation or portfolio characteristics.

Appendix Tables A1 and A2 report the corresponding decompositions for the effective number of holdings and for the equally weighted mean of the investor-level domestic portfolio-value share.

B Additional cross-tabulations and specification checks

This appendix reports additional cross-tabulations and specification checks referenced in Section 5. It covers portfolio structure by the number of distinct instruments held (Table A3); the domestic portfolio-value share by portfolio-value decile and observation date (Table A4); alternative functional forms for the relationship between portfolio value and

Table A1 Decomposition of the change in the mean effective number of holdnings

	2020 → 2021	2021 → 2022	2022 → 2023
Mean, t	3.763	4.501	4.456
Mean, t+1	4.501	4.456	4.112
Total change	0.738	-0.046	-0.344
Within stayers	0.813	0.265	0.350
Entry composition	-0.157	-0.405	-0.760
Exit composition	0.081	0.094	0.065
Stayers	9,282	17,194	26,004
Entrants	8,924	9,749	15,230
Exiters	555	1,012	939
Entrant mean	4.338	3.741	2.815
Stayer mean, t+1	4.658	4.861	4.871
Exiter mean, t	2.401	2.897	2.643
Stayer mean, t	3.845	4.596	4.521

Notes: The table applies the decomposition in Equation (8) to the equally weighted mean effective number of holdings. The entry and exit terms describe changes in the observed sample composition. Components sum to the total change up to rounding.

Table A2 Decomposition of the change in the equally Weighted mean domestic portfolio value share (pp)

	2020 → 2021	2021 → 2022	2022 → 2023
Mean, t	87.69	89.91	91.06
Mean, t+1	89.91	91.06	92.98
Total change	2.22	1.14	1.92
Within stayers	0.09	-0.15	0.12
Entry composition	1.84	1.34	1.82
Exit composition	0.29	-0.05	-0.01
Stayers	9,282	17,194	26,004
Entrants	8,924	9,749	15,230
Exiters	555	1,012	939
Entrant mean	91.83	93.42	96.08
Stayer mean, t+1	88.07	89.72	91.16
Exiter mean, t	82.82	90.74	91.46
Stayer mean, t	87.98	89.87	91.04

Notes: The table applies Equation (8) to the equally weighted mean of investor-level portfolio-value shares allocated to Romanian-classified financial instruments. It does not decompose changes in the value-weighted aggregate domestic portfolio share.

breadth (Table A5); foreign-instrument participation by first-observation cohort (Table A6); sensitivity of the headline statistics to trimming the largest accounts (Table A7); and the asset-class composition of aggregate holdings (Table A8 and Figure A1). Figures A2 and A3 provide graphical counterparts to Tables 2 and 8.

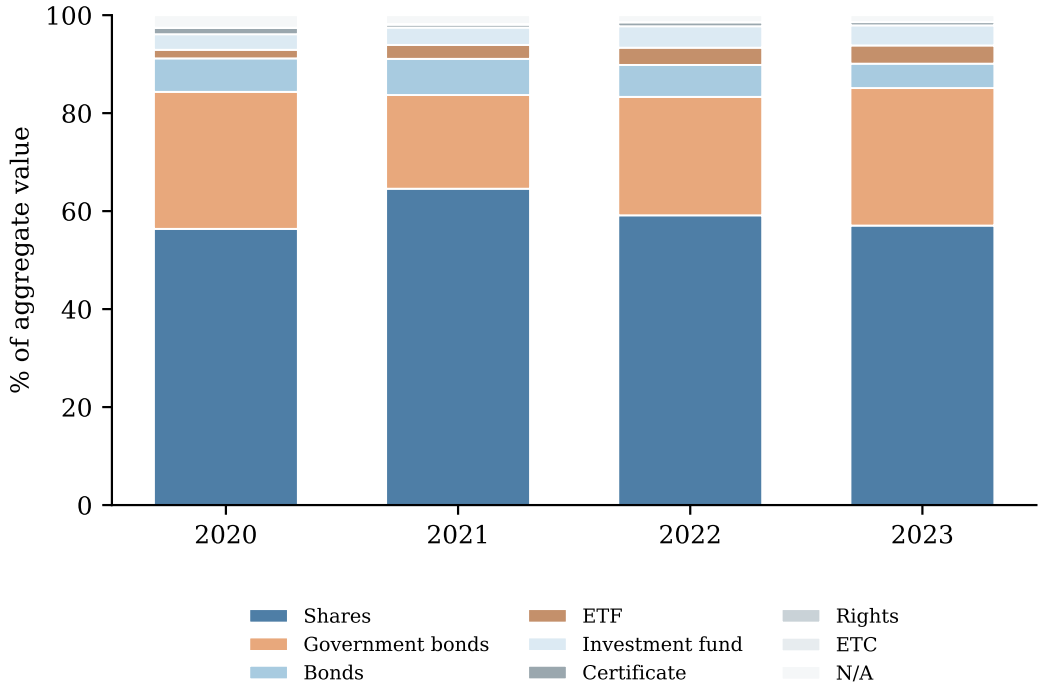


Figure A1 Asset-class composition of aggregate holdings

Table A3 Portfolio characteristics by the number of distinct instruments held

	1	2	3	4-5	6-10	11-20	>20
Investor-years	17,439	9,965	8,088	12,958	22,529	18,434	6,807
% of sample	18.1	10.4	8.4	13.5	23.4	19.2	7.1
% of value	8.0	8.1	7.7	13.0	17.7	25.1	20.4
Holds fund/ETF (%)	23.3	26.6	30.7	38.6	55.8	72.7	82.2
Holds foreign (%)	8.1	12.4	16.0	18.8	27.5	52.4	75.2
Holds gov. bond (%)	11.8	15.4	13.4	13.3	16.6	28.3	42.3
Mean N_{eff}	1	1.5	2.1	2.9	4.6	7.2	12.4
Domestic share (%)	91.9	91.1	90.7	91.6	92.1	91.7	86.7

Notes: Pooled rule-E panel. "% of value" is the share of aggregate portfolio value held by the bin. The domestic share is essentially flat across bins up to 20 instruments, while participation in funds and exchange-traded funds rises steeply. These are the two facts underlying the Hypothesis 3 and Hypothesis 4 tests.

Table A4 Domestic portfolio-value share by portfolio-value decile and observation date (%)

Portfolio-value decile	2020	2021	2022	2023
1	93.4	94.2	94.1	95.8
2	90.1	93.9	93.0	95.0
3	90.8	93.1	93.6	94.8
4	88.2	92.0	93.9	94.9
5	87.9	91.0	91.7	94.7
6	87.8	88.9	90.2	93.0
7	86.7	89.2	90.3	91.9
8	86.1	86.8	89.8	91.7
9	83.6	86.4	88.3	90.6
10	82.2	83.5	85.7	87.4

*Notes:*The table reports the equally weighted mean domestic portfolio-value share within portfolio-value deciles formed separately at each observation date. Domestic allocations are generally lower in higher portfolio-value deciles, although the relationship is not strictly monotonic within every date. The domestic share is higher in August 2023 than in December 2020 in every decile. These descriptive changes cannot be attributed exclusively to investor entry without an additional within-decile decomposition.

Table A5 Alternative functional form for or the relationship between brokerage portfolio value and breadth

	All investor-years	Deciles 1–5	Deciles 6–10
Investor-years	96,220	48,109	48,111
N on $\ln V$	0.174	0.100	0.059
N on V	0.001	0.097	0.000
$\ln N$ on $\ln V$	0.223	0.155	0.035

Notes: R^2 from bivariate regressions of the stated dependent variable on the stated regressor, pooled across observation dates. The fixed-cost mechanism implies N linear in $\ln V$; a binding minimum position size implies N proportional to V. Deciles are formed within observation date.

Table A6 Foreign-Instrument participation by entry cohort, conditional on portfolio value

	2021	2022	2023
Raw gap vs 2020 cohort (pp)	1.1	-8.3	-21.6
SE	0.7	0.7	0.6
Conditional on $\ln V$ (pp)	3.4	-3.6	-13.0
SE	0.7	0.7	0.6

Notes: Linear probability models for holding at least one foreign instrument, estimated on the 2023 cross-section with entry-cohort indicators and the 2020 cohort as the omitted category. The second pair of rows adds $\ln V_{it}$. Coefficients in percentage points, standard errors clustered by investor.

Table A7 Sensitivity to trimming the largest accounts

	Full sample	Drop top 0.1%	Drop top 1%	Drop top 5%
Investor-years	96,220	96,122	95,255	91,407
Value retained (%)	100.00	77.35	56.52	33.14
Mean N	7.97	7.96	7.88	7.53
Mean N_{eff}	4.25	4.24	4.23	4.13
Mean HHI	0.45	0.45	0.45	0.46
Domestic, VW (%)	86.86	85.17	86.53	88.21
Domestic, EW (%)	91.32	91.33	91.40	91.68
Gov. bond, VW (%)	24.95	17.76	19.93	19.68
Holds fund/ETF (%)	47.58	47.60	47.63	47.55

Notes: Accounts above the stated within-observation date percentile of portfolio value are dropped. "Value retained" shows how much of aggregate value is removed — substantial — against how little the investor-level statistics move. No "retail" cut-off is imposed anywhere in the paper; this table exists to show that the results do not depend on one.

Table A8 Asset-class composition of aggregate holdings (value shares, %)

instrument	2020	2021	2022	2023
Shares	56.38	64.57	59.14	57.05
Government bonds	27.96	19.13	24.17	28.07
Bonds	6.84	7.38	6.52	4.98
ETF	1.75	2.84	3.52	3.72
Investment fund	3.18	3.54	4.37	4.09
Certificate	1.30	0.60	0.83	0.67
Rights	0	0.12	0.01	0.02
ETC	0.03	0.00	0.01	0.00
N/A	2.55	1.81	1.42	1.39

Notes: Value-weighted shares of all client holdings. Reproduces Appendix Table A10 exactly; retained here in percentage form for reference in the text.

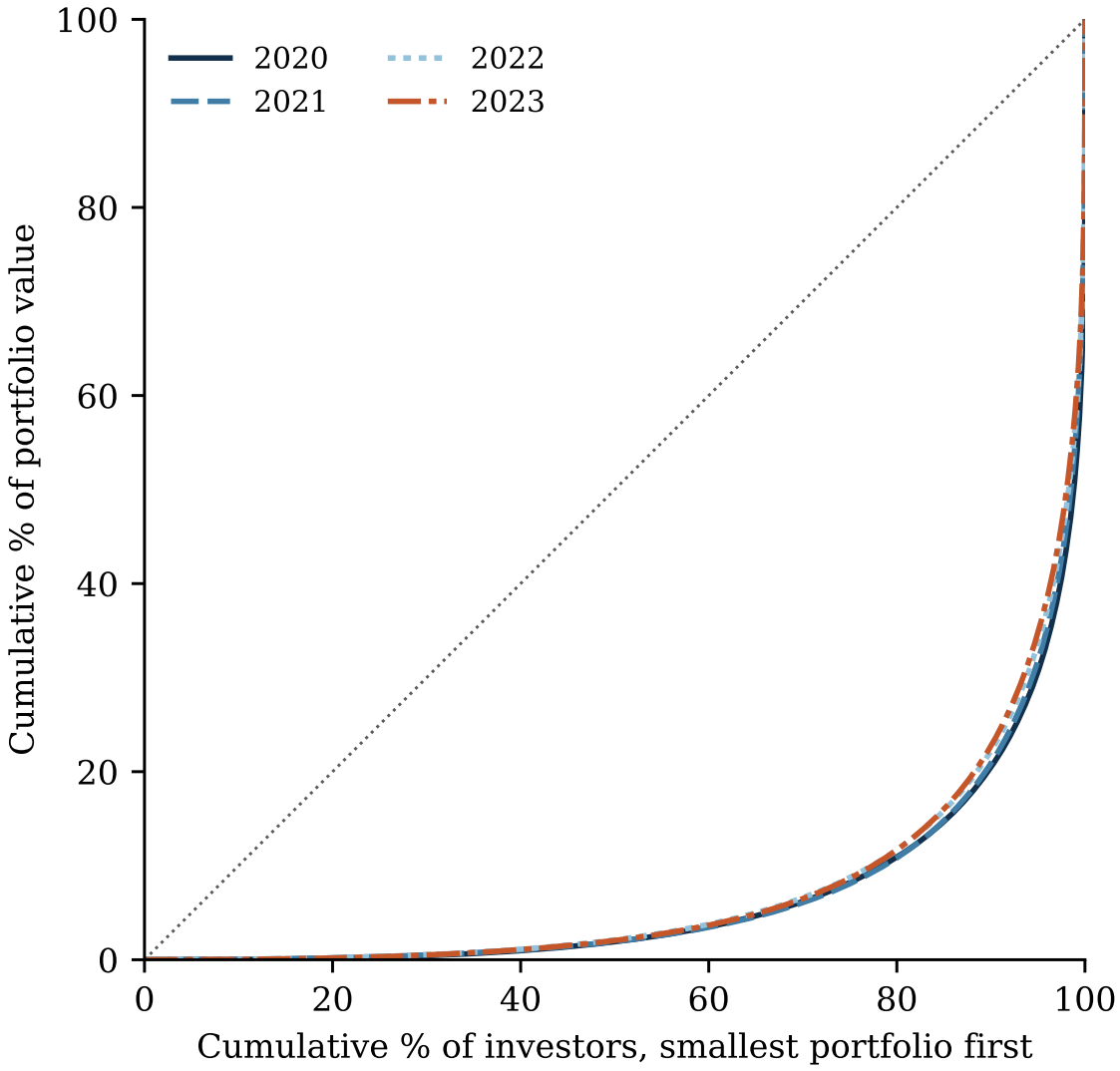


Figure A2 Concentration of portfolio value across investors

Notes: The curves describe the distribution of observed brokerage portfolio value across investors at each observation date. The Gini coefficient measures inequality in portfolio values between investors and should not be confused with the Gini coefficient of position weights within an investor's portfolio reported in Appendix Table A21. Lorenz curves of portfolio value within each observation date.

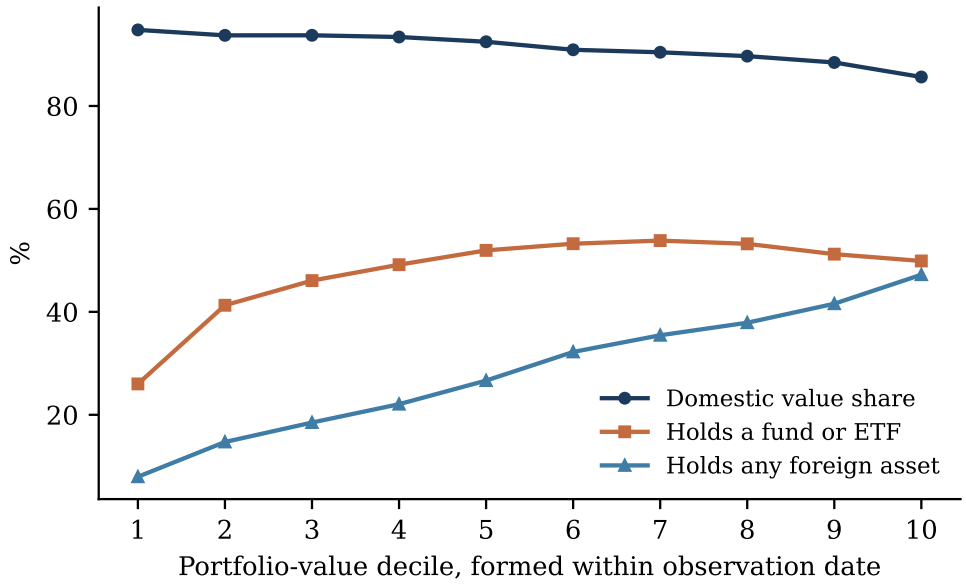


Figure A3 Domestic Portfolio Allocation, Fund/ETF Participation, and Foreign-Instrument Participation by Portfolio-Value Decile

C Composition tables

The tables in this appendix report the aggregate value of brokerage holdings by issuer country, financial-instrument type, investor characteristics, and first-observation cohort. The country classifications are subject to the limitations discussed in Section 4. Country shares refer to the classifications assigned to the financial instruments rather than the geographic distribution of their underlying economic exposures. Investor-characteristic categories include missing or unclassified observations where applicable. All portfolio values are expressed in current Romanian lei at the respective observation dates. Tables A12–A16 group aggregate brokerage holdings according to the investor characteristics available in the data. Percentages refer to shares of total brokerage portfolio value, not shares of investors. The characteristics differ in their reporting requirements and completeness, and some are self-reported rather than independently validated. Changes in the reported category shares may reflect changes in the composition of the observed investor population, the distribution of portfolio value, and the availability of investor information.

Table A9 Aggregate Brokerage Holdings by Financial-Instrument Issuer Country

Country	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
Australia	0.183	0.176	0.229	0.195	0.01	0.01	0.01	0
Austria	28.282	21.859	31.704	34.465	1.32	0.65	0.85	0.67
Belgium	1.817	0.608	0.519	0.457	0.08	0.02	0.01	0.01
Bulgaria	0.004	0.004	0.004	0.004	0	0	0	0
Canada	7.503	9.883	13.008	20.136	0.35	0.29	0.35	0.39
Cayman Islands	0.065	0.408	0.338	0.545	0	0.01	0.01	0.01
Czech Republic	0	0	0	0.114	0	0	0	0
Denmark	0.021	0.063	0.062	0.195	0	0	0	0
Finland	0.092	0.073	0.048	0.063	0	0	0	0
France	10.232	17.711	12.179	16.116	0.48	0.52	0.33	0.31
Germany	40.381	79.407	104.812	133.672	1.88	2.35	2.82	2.59
Hong Kong	0	0.738	0.856	0.837	0	0.02	0.02	0.02
India	0	0.007	0.005	0.008	0	0	0	0
Iran	0.124	0.735	0.642	0.361	0.01	0.02	0.02	0.01
Ireland	1.506	5.967	7.563	14.22	0.07	0.18	0.2	0.28
Italy	1.682	2.74	2.344	4.677	0.08	0.08	0.06	0.09
Kazakhstan	0	0.025	0.026	0.03	0	0	0	0
Luxembourg	13.329	10.745	10.402	14.891	0.62	0.32	0.28	0.29
Netherlands	16.395	23.842	28.086	36.509	0.76	0.7	0.75	0.71
Norway	0.015	0.075	0.096	0.093	0	0	0	0
Poland	2.339	0.418	0.551	0.509	0.11	0.01	0.01	0.01
Portugal	0.158	0.499	0.219	0.271	0.01	0.01	0.01	0.01
Romania	1824.142	2902.922	3269.297	4518.933	84.88	85.81	87.82	87.42
Slovakia	0.082	0.163	0.061	0.084	0	0	0	0
Spain	11.284	13.246	13.242	10.364	0.53	0.39	0.36	0.2
Sweden	1.643	4.083	3.316	4.456	0.08	0.12	0.09	0.09
Turkiye	0	0.005	0.009	0.011	0	0	0	0
United Kingdom	3.366	4.734	3.548	4.285	0.16	0.14	0.1	0.08
United States	184.404	281.738	219.493	352.76	8.58	8.33	5.9	6.82
TOTAL	2152.048	3392.065	3723.659	5169.961	100	100	100	100

Table A10 Value of holdings by financial-instrument type

Instrument type	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
Bonds	147.118	250.275	242.846	257.233	6.84	7.38	6.52	4.98
Certificate	28.041	20.422	31.036	34.682	1.3	0.6	0.83	0.67
ETC	0.724	0.166	0.201	0.165	0.03	0	0.01	0
ETF	37.667	96.454	131.094	192.394	1.75	2.84	3.52	3.72
Government bonds	601.813	648.8	900.159	1451.206	27.96	19.13	24.17	28.07
Investment fund	68.507	119.938	162.839	211.632	3.18	3.54	4.37	4.09
Rights	0	4.107	0.271	1.16	0	0.12	0.01	0.02
Shares	1213.25	2190.42	2202.319	2949.606	56.38	64.57	59.14	57.05
N/A	54.926	61.481	52.893	71.881	2.55	1.81	1.42	1.39
TOTAL	2152.046	3392.063	3723.658	5169.959	100	100	100	100

Notes: The categories reflect the financial-instrument classifications in the brokerage data. Government bonds are reported separately from other bonds. ETFs and investment funds are also reported separately and may differ substantially in their underlying exposures and investment objectives. The N/A category denotes financial instruments whose type could not be classified. Shares are calculated relative to the aggregate value of all qualifying brokerage holdings at each observation date.

Table A11 Sample holdings relative to Bucharest Stock Exchange capitalization

Year	Stocks			Bonds		
	Total mkt. cap	Our sample	Share (%)	Total mkt. cap	Our sample	Share (%)
2020	164,104.76	997.44	0.61	5,310.45	748.93	14.10
2021	248,880.31	1,846.04	0.74	8,436.68	899.08	10.66
2022	210,747.15	1,911.66	0.91	9,489.62	1,127.13	11.88
2023	309,103.11	2,518.45	0.81	12,647.15	1,676.57	13.26

Notes: The table compares the value of the sampled brokerage clients' relevant domestic stock and bond holdings with the corresponding Bucharest Stock Exchange capitalization aggregates, expressed in millions of RON. The reported ratios describe the scale of the sampled holdings relative to the exchange aggregates and should not be interpreted as the broker's share of all Romanian retail investors. The 2023 comparison is not contemporaneous: brokerage holdings are observed on August 22, 2023, whereas the exchange capitalization figures refer to year-end 2023. The relevant security universe and numerator–denominator definitions should be verified before interpreting these ratios as market shares.

Table A12 Value of holdings by sex

Sex	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
Female investors	301.281	503.171	577.431	831.79	14	14.83	15.51	16.09
Male investors	1739.907	2724.219	2917.968	4072.161	80.85	80.31	78.36	78.77
N/A	110.858	164.674	228.258	266.007	5.15	4.85	6.13	5.15
TOTAL	2152.046	3392.063	3723.658	5169.959	100	100	100	100

Notes: Values are the aggregate brokerage holdings of investors classified by sex in the investor-characteristics data. Percentages represent shares of total brokerage portfolio value, not shares of investors or average portfolio values within each category. N/A denotes missing information.

Table A13 Value of holdings by age group

Age group	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
<18	0.621	1.015	0.963	1.032	0.03	0.03	0.03	0.02
18–24	3.231	9.378	13.48	26.455	0.15	0.28	0.36	0.51
25–34	73.042	177.956	251.369	434.341	3.39	5.25	6.75	8.4
35–44	485.259	907.934	1077.854	1653.498	22.55	26.77	28.95	31.98
45–54	482.74	865.037	937.387	1322.557	22.43	25.5	25.17	25.58
55–64	262.186	432.159	442.101	614.778	12.18	12.74	11.87	11.89
65+	734.109	833.908	772.245	851.288	34.11	24.58	20.74	16.47
N/A	110.86	164.677	228.259	266.009	5.15	4.85	6.13	5.15
TOTAL	2152.046	3392.063	3723.658	5169.959	100	100	100	100

Table A14 Value of holdings by education level

Education	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
<10 years	4.526	9.489	9.657	11.166	0.21	0.28	0.26	0.22
High school	70.672	126.748	136.984	169.249	3.28	3.74	3.68	3.27
Post-secondary	37.506	53.748	50.363	57.397	1.74	1.58	1.35	1.11
University	1345.848	2020.762	2057.709	2575.791	62.54	59.57	55.26	49.82
Postgraduate	569.086	992.26	1013.19	1312.404	26.44	29.25	27.21	25.39
N/A	124.41	189.057	455.754	1043.951	5.78	5.57	12.24	20.19
TOTAL	2152.046	3392.063	3723.658	5169.959	100	100	100	100

Notes: Education categories reflect the information reported in the investor-characteristics data. The percentages describe shares of aggregate brokerage portfolio value. N/A denotes observations for which education information is unavailable. The figures should not be interpreted as population education shares or as the distribution of wealth by education among Romanian households.

Table A15 Value of holdings by income band

Income group	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
<6,000 EUR	77.183	143.04	167.534	212.106	3.59	4.22	4.5	4.1
6,000–12,000 EUR	128.864	227.292	246.56	311.357	5.99	6.7	6.62	6.02
12,000–24,000 EUR	254.985	439.241	470.405	603.692	11.85	12.95	12.63	11.68
24,000–48,000 EUR	328.112	585.7	645.542	848.904	15.25	17.27	17.34	16.42
48,000–120,000 EUR	383.458	723.736	725.574	960.125	17.82	21.34	19.49	18.57
>120,000 EUR	855.201	1083.989	1011.848	1189.109	39.74	31.96	27.17	23
N/A	124.243	189.065	456.194	1044.665	5.77	5.57	12.25	20.21
TOTAL	2152.046	3392.063	3723.658	5169.959	100	100	100	100

Notes: Income bands are based on the self-reported information recorded in the brokerage data. Percentages represent shares of aggregate brokerage portfolio value, not income shares or average portfolio values within the reported income groups. N/A denotes missing information.

Table A16 Value of holdings by reported trading experience

Experience group	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
<6 months	413.2	847.095	977.549	1313.294	19.2	24.97	26.25	25.4
6–12 months	101.274	185.936	205.604	283.574	4.71	5.48	5.52	5.49
1–3 years	130.622	247.104	284.778	379.746	6.07	7.28	7.65	7.35
>3 years	843.187	1053.017	1008.867	1126.15	39.18	31.04	27.09	21.78
N/A	663.764	1058.911	1246.859	2067.194	30.84	31.22	33.48	39.98
TOTAL	2152.046	3392.063	3723.658	5169.959	100	100	100	100

Notes: Trading-experience categories reflect self-reported information recorded in the investor profile. Percentages represent shares of aggregate brokerage portfolio value. N/A denotes unavailable experience information.

Table A17 Value of holdings by entry cohort

First year observed	Value of holdings (mln RON)				Value of holdings (% of total)			
	2020	2021	2022	2023	2020	2021	2022	2023
2020	2152.047	2914.666	2716.874	3181.048	100.00	85.93	72.96	61.53
2021	0.000	477.398	647.476	897.928	0.00	14.07	17.39	17.37
2022	0.000	0.000	359.308	594.706	0.00	0.00	9.65	11.50
2023	0.000	0.000	0.000	496.277	0.00	0.00	0.00	9.60
Total	2152.047	3392.064	3723.657	5169.958	100.00	100.00	100.00	100.00

Notes: Investors are assigned to the cohort of the first observation date in which they appear. By 2023, 38.5% of aggregate value is held by investors who entered after 2020, and 61.5% by the original cohort. These shares describe the distribution of aggregate brokerage portfolio value across first-observation cohorts. They are not a decomposition of changes in value-weighted portfolio characteristics and should not be equated with the transition-specific entry contributions in §5.6.

D Full distributions of the diversification measures

The tables in this appendix report the selected distributional statistics for portfolio breadth, concentration, and the balance of portfolio-value weights by observation date. The measures differ in their economic interpretation. Portfolio breadth counts distinct financial instruments, the Herfindahl–Hirschman index and its reciprocal summarize concentration across those instruments, and the Gini coefficient and normalized Shannon entropy describe the distribution of portfolio weights. None incorporates correlations among instrument returns or the underlying holdings of funds and ETFs.

Table A18 Number of Distinct Financial Instruments Held: Distributional Statistics

Year	N	Mean	SD	Skew.	Kurt.	Min	P10	P20	P30	P40	P50	P60	P70	P80	P90	Max
2020	9837	6.939	7.889	4.062	44.594	1	1	1	2	3	5	6	8	11	16	168
2021	18206	8.404	8.820	3.089	23.856	1	1	2	3	4	6	8	10	13	19	177
2022	26943	8.352	8.631	3.208	26.991	1	1	2	3	4	6	8	10	13	18	190
2023	41234	7.769	8.466	3.306	28.021	1	1	2	3	4	5	7	9	12	18	200
All	96220	7.968	8.535	3.292	28.022	1	1	2	3	4	5	7	9	12	18	200

Notes: The unit of observation is an investor at one observation date. N denotes the number of investor-date observations, while the remaining columns summarize the distribution of the number of distinct financial instruments held. The 2023 observation refers to August 22 rather than year-end. Percentiles describe portfolio breadth and do not account for the distribution of value across holdings or their underlying risk exposures.

Table A19 Herfindahl–Hirschman index of Portfolio Weights: Distributional Statistics

Year	N	Mean	SD	Skew.	Kurt.	Min	P10	P20	P30	P40	P50	P60	P70	P80	P90	Max
2020	9837	0.501	0.337	0.422	1.643	0.014	0.125	0.172	0.231	0.305	0.401	0.518	0.717	1.000	1.000	1.000
2021	18206	0.432	0.328	0.747	2.063	0.013	0.106	0.143	0.183	0.235	0.305	0.400	0.532	0.877	1.000	1.000
2022	26943	0.430	0.325	0.771	2.107	0.014	0.108	0.145	0.185	0.235	0.301	0.392	0.530	0.849	1.000	1.000
2023	41234	0.463	0.331	0.601	1.858	0.013	0.114	0.157	0.206	0.266	0.346	0.468	0.602	0.969	1.000	1.000
All	96220	0.452	0.330	0.655	1.924	0.013	0.111	0.152	0.197	0.255	0.334	0.441	0.574	0.935	1.000	1.000

Notes: The Herfindahl–Hirschman index is calculated as the sum of squared portfolio-value weights across the distinct financial instruments held by each investor. Higher values indicate greater concentration, and a value of one corresponds to a portfolio consisting of a single financial instrument. The pooled mean is approximately 0.45 and the median approximately 0.33. The index does not incorporate correlations among instrument returns or underlying holdings of funds and ETFs.

Table A20 Effective number of holdings: Distributional Statistics

Year	N	Mean	SD	Skew.	Kurt.	Min	P10	P20	P30	P40	P50	P60	P70	P80	P90	Max
2020	9837	3.763	3.765	3.724	34.129	1.000	1.000	1.000	1.394	1.932	2.493	3.283	4.322	5.811	8.032	69.835
2021	18206	4.501	4.128	2.882	22.479	1.000	1.000	1.141	1.878	2.502	3.276	4.262	5.458	7.011	9.394	76.011
2022	26943	4.456	3.996	2.779	20.895	1.000	1.000	1.178	1.888	2.549	3.320	4.248	5.410	6.897	9.246	73.764
2023	41234	4.112	3.863	3.015	23.749	1.000	1.000	1.032	1.662	2.136	2.888	3.759	4.865	6.353	8.777	74.603
All	96220	4.246	3.949	2.978	23.438	1.000	1.000	1.069	1.741	2.269	2.997	3.929	5.079	6.579	8.991	76.011

Notes: The effective number of holdings is calculated as the reciprocal of the investor-level Herfindahl–Hirschman index of portfolio-value weights. It represents the number of equally weighted financial instruments that would produce the observed concentration. The statistic does not account for correlations among instrument returns and should not be interpreted as the number of independent risk exposures. The mean effective number is computed by averaging investor-level reciprocals and is not the reciprocal of the mean HHI

Table A21 Gini coefficient of within-portfolio value weights: Distributional Statistics

Year	N	Mean	SD	Skew.	Kurt.	Min	P10	P20	P30	P40	P50	P60	P70	P80	P90	Max
2020	9837	0.322	0.237	0.083	2.011	0.000	0.000	0.000	0.169	0.272	0.343	0.403	0.465	0.533	0.632	0.961
2021	18206	0.344	0.225	-0.083	2.129	0.000	0.000	0.081	0.228	0.309	0.369	0.424	0.479	0.539	0.628	0.962
2022	26943	0.346	0.223	-0.109	2.149	0.000	0.000	0.095	0.236	0.312	0.373	0.427	0.479	0.539	0.626	0.966
2023	41234	0.337	0.230	-0.059	2.032	0.000	0.000	0.029	0.208	0.298	0.364	0.422	0.478	0.539	0.627	0.960
All	96220	0.339	0.228	-0.064	2.075	0.000	0.000	0.048	0.217	0.302	0.366	0.422	0.477	0.539	0.627	0.966

Notes: The Gini of weights equals zero both for a single-asset portfolio and for an equally weighted portfolio. This is why the lower percentiles are at zero and why we do not use this measure as a headline statistic. Furthermore, the range of the unnormalized Gini coefficient depends on the number of financial instruments held. Consequently, the statistic is not directly comparable across portfolios of different breadth without accounting for that dependence. The Gini coefficient measures inequality in position weights, not portfolio risk diversification.

Table A22 Normalised Shannon entropy of portfolio weights: Distributional Statistics

Year	N	Mean	SD	Skew.	Kurt.	Min	P10	P20	P30	P40	P50	P60	P70	P80	P90	Max
2020	7696	0.759	0.220	-1.454	4.679	0.000	0.426	0.619	0.719	0.784	0.831	0.868	0.899	0.929	0.964	1.000
2021	15144	0.781	0.201	-1.732	5.996	0.000	0.514	0.673	0.753	0.803	0.842	0.873	0.901	0.930	0.963	1.000
2022	22461	0.784	0.195	-1.709	5.951	0.000	0.522	0.678	0.755	0.804	0.843	0.873	0.901	0.929	0.960	1.000
2023	33479	0.772	0.203	-1.601	5.433	0.000	0.487	0.653	0.739	0.794	0.834	0.867	0.897	0.926	0.961	1.000
All	78780	0.776	0.202	-1.641	5.593	0.000	0.495	0.659	0.744	0.798	0.838	0.870	0.899	0.928	0.962	1.000

Notes: Normalized Shannon entropy is calculated as the Shannon entropy of portfolio-value weights divided by the logarithm of the number of distinct financial instruments held. The statistic is undefined for single-instrument portfolios, which are excluded from the reported distribution. Values approaching one indicate that portfolio weights are relatively balanced conditional on the number of holdings. Thus, a portfolio containing two equally weighted instruments and one containing twenty equally weighted instruments both attain a value of one. The statistic does not measure underlying risk diversification and should not be compared directly with unconditional portfolio statistics calculated using all investor-date observations.