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## Abstract

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JEL Classification: D14, D31, E44, E52, E58

Keywords: monetary policy, Household heterogeneity, Inequality, Income, wealth, Quantitative easing



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# How Does Monetary Policy Affect Income and Wealth Inequality? Evidence from Quantitative Easing in the Euro Area

April 25, 2021

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**Abstract**

This paper evaluates the impact of quantitative easing on income and wealth of individual euro area households. We first estimate the aggregate effects of a QE shock, identified by means of external instruments, in a multi-country VAR model with unemployment, wages, interest rates, house prices and stock prices. We then distribute the aggregate effects across households using a reduced-form simulation on micro data, which captures the portfolio composition, the income composition and the earnings heterogeneity channels of transmission. The earnings heterogeneity channel is important: QE compresses the income distribution since many households with lower incomes become employed. In contrast, monetary policy has only negligible effects on the Gini coefficient for wealth: while high-wealth households benefit from higher stock prices, middle-wealth households benefit from higher house prices.

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Online appendix: http://slacalek.com/research/lsMPinequality/lsMPinequality_appendix.pdf

# 1 Introduction

The collection of reliable data in recent years has allowed researchers to characterize the evolution of wealth and income distributions over time and across countries. In particular, Piketty (2013) shows that, contrary to the traditional view based on Kuznets (1955), advanced economies do not inevitably evolve toward more egalitarian societies. This fact has stimulated an intense debate about the drivers of economic inequality. In general, inequality is seen as related to the structural features of economies, such as the emergence of skill-biased technological progress (Katz and Murphy, 1992; Acemoglu, 2002; Autor, 2014), the deepening of globalization (Katz and Autor, 1999), the tendency toward the reduction in the progressivity of tax systems (Alvaredo et al., 2013) and portfolio heterogeneity (Fagereng et al., 2020; Hubmer et al., 2020).

Recently, since central banks have undertaken extensive asset purchase programmes to circumvent the lower bound on nominal interest rates, monetary policy has also been put forth as a possible driver of economic inequality (see Colciago et al., 2019, for a survey). This paper investigates how unconventional monetary policy, specifically the quantitative easing (QE) program of the European Central Bank, affects the distribution of income and wealth across individual households in the euro area. The analysis proceeds in two steps, making use of both aggregate and household-level data.

In the first stage, we estimate the transmission mechanism of a euro area QE shock. To capture the potential cross-country heterogeneity in the transmission of the common euro area monetary policy, we specify a large multi-country VAR model including macroeconomic and financial variables for the four largest countries of the euro area (France, Germany, Italy and Spain). For each country we include, among others, the variables related to the dynamics of household income and wealth: the unemployment rate, wages and house prices.[1]

We identify the QE shock by means of an external instrument approach (see Stock, 2008; Stock and Watson, 2012; Mertens and Ravn, 2013; Ramey, 2016; Miranda-Agrippino and Ricco, 2019). Gertler and Karadi (2015) suggests that the high frequency changes in financial variables recorded during the policy announcements of central banks could be used as external instruments to identify monetary policy shocks. We follow this insight and construct the external instrument for the QE shock exploiting the term-structure of overnight indexed swap rates (OIS). In order to exclusively capture the effects of QE, we use as our external instrument the QE factor of Swanson (2021) and Altavilla et al. (2019), an aggregate of the changes in the yield structure of the OIS rates recorded during the ECB press conference, which is orthogonal to two additional policy factors capturing forward guidance and conventional monetary policy and is constrained to explain a negligible share of the volatility in OIS rates in the period preceding the Lehman collapse, when QE-type policies were not in place.

[1] We also include GDP and the GDP deflator for each country, and long-term interest rates and stock prices for the euro area. We do not impose any restrictions on the dynamic relationships across variables, as for example in the panel VAR literature. The large dimension of the model (22 variables, in (log-)levels, five lags) is handled using Bayesian estimation methods with informative priors which, as suggested by De Mol et al. (2008) and Bańbura et al. (2010), controls for overfitting while at the same time extracting the valuable information in the sample. The informativeness of the prior distributions is set according to the hierarchical BVAR procedure developed in Giannone et al. (2015).

Our first result is that allowing for cross-country heterogeneity in the transmission mechanism is important, as the impulse responses of unemployment rates and asset prices vary across countries: for example, the unemployment rate in Spain responds considerably more to the QE shock than in Germany. Our results are broadly in line with those of previous studies of the effects of central bank asset purchases, which also find that asset purchase programs such as QE have noticeable effects on the real economy (for an extensive recent survey, see Dell'Ariccia et al., 2018, and our online Appendix C).[2]

However, aggregate cross-country heterogeneity is only one of the possible relevant dimensions to capture the different impact of QE across households. For example, the QE shock may result in heterogeneous impacts on households also because of the substantial differences in their sources of income (e.g., employment status, labor vs financial income) and their portfolio holdings (holdings of real estate, shares and bonds). Consequently, in the second stage, we use simulation techniques to distribute the aggregate effects estimated in the VAR across the individual households using data on their asset and income composition. The analysis relies on the Household Finance and Consumption Survey (HFCS), a dataset which collects detailed household-level information on balance sheets, income and socio-demographic variables for European countries (similar to the Survey of Consumer Finances for the US).

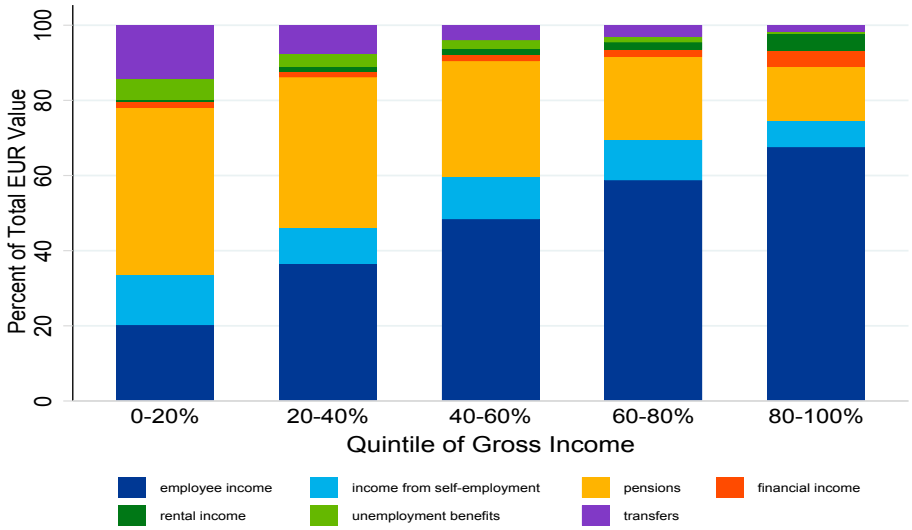
Our analysis captures the transmission of QE to households via three channels: (i) income composition, (ii) portfolio composition and (iii) earnings heterogeneity. The two composition channels operate via the heterogeneous reaction of various income and wealth components to monetary policy. Figure 1 reports the sources of income for the aggregate population of France, Germany, Italy and Spain. Both the level and the share of key income components vary substantially with the level of household income. In particular, the households in the lowest income quintile earn only roughly 20% of their gross income as employee income, while those in the top quintile about 60%. Similarly, the share of financial and rental income increases from 2% to almost 10%. In contrast, the share of transfers and unemployment benefits declines across income quintiles from almost 20% to about 3%.

Figure 2 shows that the composition of household wealth is similarly varied. In particular, the share of self-employment business wealth (private businesses) and stock market wealth (shares) on total assets in the top net wealth quintile is substantially larger, while the share of real estate is lower. To empirically capture the two composition channels, we update the components of income and wealth at the household level using the aggregate impulse responses for wages and for house, stock and bond prices.[3] The earnings heterogeneity channel, instead, relates to the heterogeneous reaction of the employment status to monetary policy. To capture this channel, we run a reduced-form

[2]Our results, both in terms of the aggregate impulse responses of the multi-country VAR and the implications for individual households remain basically unaffected if we rely on a different identification scheme based on sign restrictions, whereby an expansionary QE shock is identified by assuming that it has a positive effect on GDP in all countries. The results based on this alternative identification scheme are available in the ECB Working Paper draft, Lenza and Slacalek (2018).

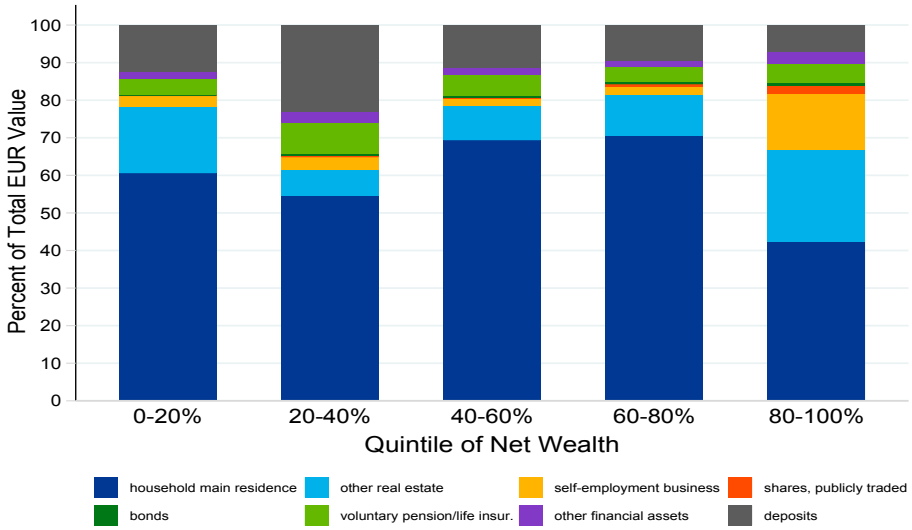
[3]In the baseline setup we assume that household portfolios are not rebalanced in response to the QE shock. This assumption is supported by the empirical evidence on considerable inertia in household portfolios, e.g., Ameriks and Zeldes (2004), Brunnermeier and Nagel (2008), Andersen et al. (2020) and others. We relax this assumption in one of our robustness simulations.

**Figure 1** Composition of Income




**Source**: Household Finance and Consumption Survey, wave 2014
**Note**: The figure shows how the share of income components in total gross income varies across quintiles of gross income. Unemployment benefits and transfers include regular social transfers (except pensions) and private transfers. The figure shows an aggregate of France, Germany, Italy and Spain.

**Figure 2** Composition of Total Assets




**Source**: Household Finance and Consumption Survey, wave 2014
**Note**: The figure shows how the share of components in total assets varies across quintiles of net wealth. Other financial assets include managed accounts, mutual funds and money owed to households. The figure shows an aggregate of France, Germany, Italy and Spain.

simulation which redistributes the aggregate decline in unemployment across individuals depending on their demographic characteristics: some unemployed individuals become employed and receive a substantial increase in (labor) income, as they start earning wages rather than unemployment benefits. The simulation ensures that the reduction of the unemployment rate in the household data is consistent with the aggregate drop in unemployment in the VAR impulse responses.

Our empirical results show that accounting for household heterogeneity in income and wealth is important for describing the effects of quantitative easing on income and wealth inequality. For income, the overall effect of quantitative easing is dominated by the earnings heterogeneity channel: transitions from unemployment to employment account for almost all of the increase in income across households. Importantly, the contribution of this channel is particularly strong in the lower segment of the income distribution. For households in the bottom income quintile, one year after the realization of an exogenous QE shock driving down the term spread by 30 basis points, the unemployment rate declines by 2 percentage points and mean income increases by almost 2%. Due to the relatively muted response of wages to the shock, the effect on wages of all existing workers, the intensive margin, is quite small.[4]

Overall, QE reduces income inequality because the earnings heterogeneity channel dominates the income composition channel, which increases more incomes at the top: the Gini coefficient for gross household income declines from 43.16% to 43.03%. While the effects are likely to fade away over longer horizons, given the likely transient nature of the effects of monetary policy, this evidence suggests that quantitative easing contributes to support vulnerable households. Our main robustness checks pertain to alternative scenarios in which financial income strongly increases due to QE. While the increase in financial income is particularly beneficial for the top tail of the income distribution, its contribution to the changes in total income is limited and, hence, it does not significantly change our results on income inequality.

We then investigate how QE changes the wealth distribution via the portfolio composition channel. The policy increases the value of stocks, mostly held by wealthier households. This effect, by itself, would lead to an increase in wealth inequality. However we find that Gini index of net wealth in fact remains basically unaffected by the QE shock, increasing negligibly from 69.17% to 69.18% because the effects of the increase in stock prices on net wealth are essentially offset by those related to housing wealth. This reflects the fact presented in Figure 2 that a large share of the population, i.e., about 60% of euro area households, own their main residence and that this asset category has a larger weight for the mid- and low-quintiles of the distribution. The conclusion that the Gini index for the wealth distribution remains largely unaffected is robust to allowing for some rebalancing of financial portfolios and more differentiated responses of house prices to quantitative easing.

Our paper is related to the growing literature on the effects of monetary policy on inequality. Coibion et al. (2017) use quarterly data from the US Consumer Expenditure

[4]The calibration of the size of the QE shock to a 30 basis points drop in the term spread is close to the lower boundary estimated for the effect of the first QE announcement in the euro area, see for example Altavilla et al. (2015).

Survey in a VAR with narrative shocks to estimate the effects of standard monetary policy on the Gini coefficients for consumption and income. A few papers follow the approach of Coibion et al. (2017) to assess the impact of standard policy on inequality in other countries, notably Mumtaz and Theophilopoulou (2017) for the UK, Guerello (2018) for the euro area and Furceri et al. (2018) in a panel data study of 32 advanced and emerging market countries. We focus on unconventional monetary policy, specifically the euro area QE, and we assess the effects of monetary policy both on income and wealth inequality. By also looking at wealth inequality, we contribute to the debate on the relative importance of direct and indirect effects of monetary policy on consumption, since such effects can be estimated only by considering the transmission channels involving both income and wealth (Kaplan et al., 2018; Auclert, 2019; Slacalek et al., 2020). Moreover, our approach combines analysis based on macro and micro data to capture and assess the relative importance of the different steps of the transmission of the QE shock to the inequality indices, with the aim to inform economic modelling and policy decisions. Casiraghi et al. (2018) (on Italian data) and, at least partly, Bunn et al. (2018) (on UK data) also focus on unconventional monetary policy.[5] We precisely identify the effects of quantitative easing in a multi-country VAR for four euro area countries which, among other things, also accounts for the cross-country spillovers of monetary policy impulse. In addition, our approach to distribute the aggregate impulse responses of income components accounts for the transitions from unemployment to employment (the extensive margin). Adam and Tzamourani (2016) quantify the effects of hypothetical scenarios on the evolution of various asset prices (stock, bond and house prices) focusing exclusively on the wealth of euro area households. Our analysis has a different focus from the work of Kuhn et al. (2020), who describe the *unconditional* historical evolution of the US wealth distribution, highlighting the contribution of house prices for the lower 90% of the households and of stock prices for the top 10% (see Martínez-Toledano, 2020, for corresponding analysis for Spain). Our purpose, instead, is to isolate the effects of quantitative easing on inequality and, for this reason, we use impulse responses from a VAR to identify the changes in the wealth distribution *conditional* on the effects of quantitative easing.

The remainder of the paper is organized as follows. Section 2 outlines our empirical method based on a multi-country VAR model and a simulation on household-level income and wealth data. Section 3 describes and interprets the empirical results and the main robustness checks. Section 4 concludes.

# 2 Empirical Methodology

We estimate the effects of QE on wealth and income of individual households in two steps: First, we estimate a Bayesian VAR model on aggregate data and identify the effects of monetary policy shocks at the aggregate level. Second, we undertake a reduced-form

[5] For the analysis of the macroeconomic effects of the policies, Casiraghi et al. (2018) relies on the Banca d'Italia's assessment of two different "unconventional monetary policy" scenarios, with and without financial stress in the economy. Bunn et al. (2018), instead, is based on the evaluation of the effects of the conventional and unconventional monetary policy stimulus of Bank of England in 2008–2014 reported in Carney (2016).

simulation using micro data to distribute the aggregate effects across the components of income and wealth across individual households. This section describes both steps in detail.

## 2.1 The BVAR Model and the Identification of the QE shock

We identify the effects of QE using a large multi-country vector autoregression (VAR).[6] Such setup allows us to estimate possibly heterogeneous country responses to a common euro area QE shock. In more detail, to capture the dynamic interrelationships among the variables, we adopt the following VAR setting:

$$\begin{aligned} y_t &= C + B_1 y_{t-1} + \cdots + B_p y_{t-p} + \epsilon_t, \\ \epsilon_t &\sim \mathcal{N}(0, \Sigma), \end{aligned}$$

where $y_t$ is an $N$-dimensional vector of time-series, $B_1, \ldots, B_p$ are $N \times N$ matrices of coefficients on the $p$ lags of the variables, $C$ is an $N$-dimensional vector of constants and $\Sigma$ is the covariance matrix of the errors. The model is specified in terms of the annualized (log-)levels of the variables and, in our specification, we have $N = 22$ and $p = 5$. In particular, for France, Germany, Italy and Spain, we consider real GDP, the GDP deflator, the unemployment rate, house prices and wages. We also include long-term interest rates and stock prices for the euro area. The variables are available at the quarterly frequency, for the sample 1999Q1 to 2016Q4.

Potentially, this model may be subject to the "curse of dimensionality" due to the large number of parameters to be estimated, relative to the available sample. In such circumstances, the estimation via classical techniques would very likely result in overfitting the data and large estimation uncertainty. De Mol et al. (2008) and Bańbura et al. (2010) showed that imposing informative priors which push the parameter values of the model toward those of naïve representations (such as, for example, the random walk model) reduces estimation uncertainty without introducing substantial bias in the estimates, thanks to the tendency for most macroeconomic and financial variables to co-move. In fact, in presence of comovement, the information in the data strongly "conjures" against the prior, so that the parameter estimates reflect sample information even if very tight prior beliefs are enforced.

For this reason, we estimate the model with Bayesian techniques. The prior for the covariance matrix of the residuals $\Sigma$ is Inverse Wishart, while the prior for the autoregressive coefficients is normal (conditional on $\Sigma$). As it is standard in the BVAR literature, we follow Litterman (1979) and parameterize the prior distribution to shrink the parameters toward those of the naïve and parsimonious random walk with drift model, $y_{i,t} = \delta_i + y_{i,t-1} + e_{i,t}$. Moreover, we also impose a prior on the sum of the VAR coefficients to address the issues raised by the tendency of VAR models to overfit the data via their deterministic component (see Sims, 1996, 2000; Giannone et al., 2019, for an extensive discussion of this pathology of VARs). The full specification and the estimation method used for the VAR model follows Giannone et al. (2015). The setting of the

[6] See Appendix B for the details on the macroeconomic database.

prior distributions depends on the hyperparameters which describe their informativeness for the model coefficients. For these parameters, we follow the theoretically grounded approach proposed by Giannone et al. (2015), which suggests to treat them as random variables, in the spirit of hierarchical modelling, and conduct posterior inference also on them. As hyper-priors (i.e., prior distributions for the hyperparameters), we use proper but almost flat distributions. For details on specification of the prior distribution see Appendix A.[7]

To estimate the effects of quantitative easing, we identify a QE shock by means of an external instrument approach. Here we provide the intuition for this method, for an extensive and rigorous treatment, see Stock (2008); Stock and Watson (2012); Mertens and Ravn (2013); Ramey (2016); Miranda-Agrippino and Ricco (2019).[8] Define the moving average representation of the VAR above as:

$$y_t = \sum_{k=0}^{\infty} D_k \epsilon_{t-k}.$$

The $N$-dimensional vector of structural shocks $\varepsilon_t$ is linearly related to the vector of the VAR reduced form residuals via the $N$-dimensional square matrix $\Theta_0$:

$$\epsilon_t = \Theta_0 \varepsilon_t.$$

Let us also assume, without loss of generality, that the QE shock is ordered first in the vector of structural shocks and it is defined as $\varepsilon_{1,t}$. Once the first column of $\Theta_0$, denoted $\Theta_{0,1}$, is retrieved, the moving average VAR representation can be used to find the impulse response of each variable $y_t$ to the shock $\varepsilon_{1,t}$. An external instrument $z_t$ for the structural shock $\varepsilon_{1,t}$ is essentially a variable that is correlated with that structural shock and uncorrelated with all the other $N-1$ structural VAR shocks:

$$\begin{aligned} \mathbf{E}(z_t \varepsilon_{1,t}) &= \zeta, \\ \mathbf{E}(z_t \varepsilon_{j,t}) &= 0, j = 2, \ldots, N. \end{aligned}$$

Then the covariance between $z_t$ and the reduced form VAR shocks is:

$$\mathbf{E}(z_t \epsilon_t) = \zeta \Theta_{0,1},$$

which can be used to identify $\Theta_{0,1}$ up to a scaling constant.

Of course, the method relies on the existence of a suitable instrument to identify the QE shock. To address this potential challenge, we follow the insight of Gertler and Karadi (2015), which suggests that the high frequency changes recorded in specific financial variables during the policy announcements of central banks could be used as external instruments to identify monetary policy shocks. The idea is that such changes in financial variables are correlated to the monetary policy shocks and, at the same

[7] A few papers lend support to this strategy to model cross-country macroeconomic data, showing that VAR models of the type we adopt in this paper provide accurate out-of-sample forecasts of macroeconomic and financial variables in the euro area (see, for example, Angelini et al., 2019; Capolongo and Pacella, 2019). A similar framework has been also used to estimate the effects of common euro area monetary policy shocks on various countries by Altavilla et al. (2016) (for both standard monetary policy and outright monetary transactions, OMT) and Mandler et al. (2016) (for standard monetary policy shocks). To appropriately capture the transmission channels of QE to different components of household wealth and income, we add more variables such as house prices to the existing frameworks.

[8] In this paper we use the toolbox developed by Miranda-Agrippino and Ricco (2019).

time, they are unlikely to reflect other sources of shocks given that the monetary policy announcements are the main drivers of the surprises in financial variables over narrow time windows around those announcements.

To derive a specific external instrument for the euro area QE shock, we use the changes in the OIS rates with maturity from one month to ten years recorded during the Eurosystem press conferences (available in the Monetary Policy Database of Altavilla et al., 2019) in which the ECB President announces and describes the monetary policy decisions taken by the Governing Council.[9] Gürkaynak et al. (2005) pointed out that the changes in financial variables during policy announcements are likely to reflect more than one type of monetary policy measure, especially after the collapse of Lehman Brothers paved the way for unconventional monetary policy measures. Hence, we take additional steps to disentangle the fluctuations in OIS rates due to QE from those due to other policy announcements. Specifically, we use as external instrument the so called QE factor of Swanson (2021) and Altavilla et al. (2019), an aggregate of the changes in the yield structure of the OIS rates during the ECB press conferences. The QE factor is identified by assuming (i) that it is orthogonal to the two policy factors capturing forward guidance and conventional monetary policy and (ii) that it explains only a negligible share of the volatility in the OIS rates during the press conferences preceding the Lehman crisis, when QE-type policies were not in place. Altavilla et al. (2019) shows that the QE factor does not explain much of the volatility in the short-term segment of the OIS yield curve, while it is a relevant driver of the long-term segment, lending support to the idea that the factor correctly captures the fluctuations in OIS rates due to QE.

As mentioned above, the external instrument approach identifies the QE shock up to a scaling constant. To pin down the constant to a reasonable value, we set the size of the shock to imply a 30 basis points impact reduction in the euro area long-term interest rate, the lower boundary of the estimated effects of the first QE announcement on the euro area long-term bond yields (Altavilla et al., 2015).

## 2.2 The Reduced-Form Simulation on Household-Level Wealth and Income Data

We distribute the aggregate impulse responses of unemployment, wages and asset prices across individual households depending on the structure of their assets and sources of their income. Table 1 provides an overview of the methodology.

We use the 2014 wave of the Household Finance and Consumption Survey (HFCS). The HFCS is a unique ex ante comparable household-level dataset, which contains rich information on the structure of income and household balance sheets and their variation across individual households. The dataset also collects information about socio-demographic variables, assets, liabilities, income and indicators of consumption. The reference year, 2014, matches quite well the start of the Asset Purchase Programmes.

[9] The quantitative easing program of the ECB is defined as Asset Purchase Programme (APP). It started in January 2015 to address the risks of a long period of low inflation. The APP includes various purchase programmes under which private sector securities and public sector securities (including sovereign bonds) are bought. For an early assessment of the macroeconomic effects of the APP see Andrade et al. (2016).

**Table 1** Modeling of Responses of Wealth and Income Components at Household Level

| Wealth / Income Component | Modeling Procedure |
|---|---|
| **Real Assets** | |
| Household's main residence | Multiplied with response of house prices (robustness: heterogeneity in house prices) |
| Other real estate property | Multiplied with response of house prices (robustness: heterogeneity in house prices) |
| Self-employment businesses | No adjustment |
| **Financial Assets** | |
| Shares, publicly traded | Multiplied with response of stock prices (in the baseline; robustness: some trading) |
| Bonds | Multiplied with response of bond prices (based on long-term rate) |
| Voluntary pension/whole life insurance | No adjustment |
| Deposits | No adjustment |
| Other financial assets | No adjustment |
| **Debt** | |
| Total liabilities (mortgage + non-mortgage debt) | No adjustment |
| **Gross Income** | |
| Employee income | Extensive margin: Some unemployed become employed are receive wage<br>Intensive margin: Multiplied with response of wages (compensation per employee) |
| Self-employment income | Multiplied with response of wages (compensation per employee) |
| Income from pensions | No adjustment |
| Rental income from real estate property | No adjustment |
| Income from financial investments | No adjustment (in the baseline; robustness: grows by 2.1% or country-specific) |
| Unemployment benefits and transfers | If becomes employed, replace with wage (otherwise no adjustment) |

We focus on the four largest euro area countries, in which the (net) sample ranges roughly between 4,500 households (Germany) and 12,000 households (France).[10] To adequately capture the top tail of the distribution, wealthy households in Spain, France and Germany are over-sampled.

### *2.2.1 Estimating the Effects of Quantitative Easing on Household Income: The Earnings Heterogeneity and the Income Composition Channels*

Starting with our baseline characterization of the income composition channel, which in subsequent discussion we will also refer to as the intensive margin of QE, Figure 1 shows that the key income component for most households is income from employment and self-employment. We use impulse responses of wages to assess how these income components are affected by QE at the household level. For income from rental of properties, financial investments and pensions, instead, we assume that there is no change due to QE.[11] In section 3.2.3 we provide a robustness analysis to gauge the relevance of this no-change assumption for some categories of income, such as financial income.

The earnings heterogeneity channel pertains to the effect of monetary policy on employment. We model this extensive margin as follows. The aggregate results imply that quantitative easing reduces the aggregate unemployment rate. In turn, micro data on employment and income can be used to simulate which unemployed people become employed and by how much their incomes increase. The simulation, which broadly follows the setup of Ampudia et al. (2016), is divided in two steps and runs at the individual level (not at the household level); the results are then aggregated to household level.

**Step 1: Probit Simulation for the Employment Status**
In the first step, we distribute the aggregate decline in unemployment across individuals, using a probit regression which takes into account individual characteristics. This allows us to pin-down which individuals become employed as a result of QE.

For each country $c$, we first estimate a probit model regressing individual's $k$ employment status $S$ on her demographic characteristics:

$$\Pr(S_k = 1 | V_k = v_k) = \Phi(v_k' \beta_c), \tag{1}$$

where $V$ denotes demographics: gender, education, age, marital status and the number of children; $\Phi(\cdot)$ denotes the normal cdf. For each individual we denote the fitted values, the estimated probability of being employed, as $\hat{S}_{c,k}$ and use it to simulate who becomes employed as a result of QE. For each person $k$ we then draw a uniformly distributed random 'employment' shock $\xi_k$. If the value of $\xi_k$ is sufficiently below

[10] See Household Finance and Consumption Network (2016), in particular Table 1.1, for information on the 2014 wave of the HFCS.

[11] Regarding rental income, we assume it does not respond to monetary policy easing. In the long-run rents are positively linked with house prices, so that increases in house prices tend to co-move with increases in rents. However, when focussing specifically on the response to monetary policy shocks, recent estimates (e.g., Dias and Duarte (2019) for the US and Corsetti et al. (2020) for the euro area) find a negative relationship between rents and house prices via the following mechanism: Monetary policy easing reduces mortgage rates and makes homeownership more attractive. Consequently, renters tend to buy housing and rents decline.

$\hat{S}_{c,k}$ and the person is actually unemployed, she becomes employed. This implies that people with higher $\hat{S}_{c,k}$ are more likely to become employed (although even people with lower $\hat{S}_{c,k}$ can become employed if they draw a low $\xi_k$). The threshold for $\xi_k - \hat{S}_{c,k}$, which determines how many people become employed, is set so that the number of newly employed individuals is for each country consistent with the aggregate decline in unemployment, as estimated in the VAR impulse response.[12] We repeat the simulation many times and report the average results across repetitions.[13]

**Step 2: Imputation of Labor Income**
In the second step we replace unemployment benefits of people who are newly employed with wage, which is estimated based on their demographic characteristics. Specifically, the log of wage of newly employed individuals is estimated by a two-step Heckman model of the system of wage and selection equations. Our exclusion restrictions are the marital status and the presence of children. We assume that these factors may affect the work status but not the wage of the employed. The remaining regressors in the model are gender, education and age.[14]

The estimates from the Heckman selection model are in line with the evidence from the literature for standard market economies and with our prior expectations (see online Appendix C for details). Wages increase with age and education, and are higher for men. As for the selection equation, the effects of the presence of children of being married are both positive. The probability of the selection into employment also increases with age and education, and is higher for men in Italy and Spain (and insignificant in France and Germany). The implied average replacement rate for after-tax labor income of the newly employed is roughly 45–50%.

### *2.2.2 Estimating the Effects of Quantitative Easing on Household Wealth: The Portfolio Composition Channel*

To simulate the effects of quantitative easing on wealth, i.e., to capture the portfolio composition channel, we use the detailed *quantitative* information about holdings of various asset classes by each household in the HFCS (i.e., we know the nominal market value of each asset class owned by households). The effects of monetary policy on household wealth are obtained by multiplying the holding of each asset class (in EUR) by the corresponding change in asset prices given by the VAR impulse response.

In particular, our VAR includes three asset price variables: house prices, stock prices and bond prices. We multiply the holdings of housing wealth—i.e., household's main residence and other real estate—by house prices. We multiply the holdings of shares by stock prices. We assume that the value of self-employment businesses is unaffected by the QE shock, due to the difficulty to reliably measure the value of such component of

[12] In practice, we sort unemployed individuals by their value of $(\xi_k - \hat{S}_{c,k})$ and those with the lowest rank become employed until the reduction in the unemployment rate matches the value given by the impulse response. We use survey weights in this calculation.

[13] The empirical results in the paper are based on 500 replications.

[14] We chose the classic Heckman procedure to impute labor income when it is not observed. Alternative imputation methods based on matching techniques could be considered; see, e.g., Olivetti and Petrongolo (2008) and many others.

wealth.[15] Finally, we multiply the holdings of bonds by the change in the price of the 10-year bond implied by the decline in the long-term rate.

This calculation assumes that households do not adjust their portfolios in response to monetary policy. The assumption of no rebalancing seems a reasonable first-order approximation for two reasons. First, we consider responses to relatively *small* monetary policy shock over the short-run horizon of several quarters. Related, the literature documenting the effects of quantitative easing on portfolio rebalancing (Koijen et al., 2017) finds rather small effects for the household sector (in the aggregate data). Second, substantial micro evidence exists on the sluggishness in household portfolios. This holds not only for very illiquid assets (such as housing) but also for many financial assets. For example, a well-known paper by Ameriks and Zeldes (2004) documents that almost half of the households in their data on retirement accounts (held by TIAA–CREF) made no active changes to their portfolio of stock over the *nine-year* period they consider. Similar findings are reported in Bilias et al. (2010): The bulk of US households exhibit considerable inertia in their stock portfolios (held in brokerage accounts). Several papers examine inertia in household portfolios using high-quality administrative data. Fagereng et al. (2021) document evidence on the limited extent of rebalancing of illiquid and risky assets in response to receiving a lottery prize in Norwegian data. Using Danish data, Andersen et al. (2020) study the substantial inaction of households regarding mortgage refinancing. In Swedish data, Calvet et al. (2009) find very weak active rebalancing in the household sector as a whole, though at the household-level active rebalancing compensates about half of idiosyncratic passive variations in the risky share and is stronger for financially sophisticated households. In section 3.2.3 below, we investigate how robust the results are to assuming some rebalancing in holdings of stocks and bonds, also accounting for more reallocation by wealthy households.

# 3 Empirical Results

This section describes our estimates, first focusing on the effects of monetary policy on aggregate variables identified using the VAR model, then considering the effects on wealth and income of individual households via the three channels described in the previous section: (i) income composition, (ii) portfolio composition and (iii) earnings heterogeneity.[16]

[15] As described in Table 1, we assume that other classes of net wealth, most importantly deposits and liabilities remain unaffected by monetary policy. For the time period we focus on—since 2014—this seems reasonable as the short-run interest rate was at the zero lower bound. The HFCS also records holdings of voluntary pensions, for which we in the baseline scenario assume they are unaffected by stock prices. Data on Euro area insurance corporation and pension fund statistics, `https://www.ecb.europa.eu/press/pr/stats/icpf/html/index.en.html`, indicate that pension funds hold a small fraction of their assets in stocks, i.e., about 9% of total assets is held in equities (2016Q4). Notice however that 21.5% is held in investment funds, for which it is difficult to determine what fraction of their assets they hold in stocks.

[16] We do not consider other channels of transmission, such as the interest rate exposure channel of Auclert (2019) and the inflation channel of Doepke and Schneider (2006). These channels are analyzed quantitatively in Slacalek et al. (2020).

## 3.1 Aggregate Effects of Quantitative Easing

This section reports the aggregate responses of the variables included in our multi-country VAR to an QE shock. The shock is meant to be expansionary and implies a 30 basis point drop in the euro area ten-year government bond yields. The full set of impulse responses is reported in the online Appendix C. The impulse responses for GDP and unemployment are qualitatively in line with the previous literature, which also finds relevant and statistically significant effects of asset purchases—see Dell'Ariccia et al. (2018) (e.g., their Table 1) for an up-to-date overview of the literature. We also find that QE stimulates asset prices and nominal variables such as the GDP deflator and wages, although the latter with a more marked heterogeneity across countries. In Figure 3, we focus on the median impulse responses of the variables that play an important role in the subsequent analysis on individual households.

The shock has a relatively short-lived impact on the long-term bond yields, whose median response is close to zero already after four quarters. The peak response of stock prices is quite large—15% at the peak after three quarters—but fades away substantially thereafter. The country-specific impulse responses in Figure 3 document the extent of heterogeneity across the four countries. House prices increase in all countries except Spain (where their response is not significant and even mildly negative at the median), with the peak responses ranging around 0.5–2%. The responses of the labor market variables also display cross-country heterogeneity. For what concerns the unemployment rates, the heterogeneity is more moderate and mostly appears in the size of the median impact, with a drop in all countries and a stronger effect in Spain. The response of wages is instead rather muted but also varies in sign, with some decreases in Spain and Germany and increases in the other countries; this marked cross-country heterogeneity of the response of wages to common shocks across the euro area countries is a known feature also documented in Angelini et al. (2019).[17]

## 3.2 Effects of Quantitative Easing on Individual Households

We report the estimates of the effects on income and wealth of individual households using a series of figures with 'micro' impulse responses implied by the micro-simulation described in section 2.2. The impulse responses are grouped in terms of quintiles of the income and wealth distributions.

[17] In our previous draft, available in the ECB working paper series Lenza and Slacalek (2018), we identify the QE shock by means of sign restrictions, similarly to Baumeister and Benati (2013). Specifically, we identify the effects of asset purchases using a combination of zero and sign restrictions (employing the algorithm of Arias et al., 2018). The main identifying assumption there is that an expansionary asset purchase shock decreases the term spread (defined as long-term minus short-term interest rate, where the short-term rate is the 3-month Euribor and the long-term rate is the euro area 10-year government benchmark bond yield) and has a positive impact on real GDP in the four countries under analysis. The responses of all other variables, notably the GDP deflator, the unemployment rate, wages and house prices in the four countries and stock prices, are left unrestricted. The effects of a QE shock identified by means of sign restrictions are qualitatively and quantitatively comparable with those we obtain by means of external instruments. The most relevant exception is that the response of stock prices is more persistently positive, when the shock is identified by means of external instruments.

**Figure 3** Quantitative Easing Scenario: VAR Impulse Responses


**Note**: The figures plot median responses to the asset purchase scenario. House prices: percentage deviation from baseline levels; wages: percentage deviation from baseline levels; unemployment rate: deviation from baseline level; stock price: percentage deviation from baseline levels.

### *3.2.1 Effects on Household Income: The Earnings Heterogeneity and the Income Composition Channels*

In the baseline setup, the effects of QE on income arise via two channels: (i) the earnings heterogeneity—the increase in income as unemployed people become employed (the extensive margin) and (ii) the income composition channel—the effect on labor income for all employed people due to the change in wages (the intensive margin).

Let us first investigate the earnings heterogeneity channel in isolation. Figure 4 shows the impulse responses of the unemployment rate by (country-level) income quintiles. Notice that the effect of an expansionary QE shock on the distribution of unemployment is not clear, a priori, because there are two countervailing factors that can affect the response of unemployment across income quintiles. On the one hand, higher income individuals have generally more favourable demographics (for example, a higher level of education) and, hence, also a higher estimated probability to become employed in response to the expansionary QE shock.[18] On the other hand, the bottom right panel of Figure 4 shows that the number of unemployed is heavily skewed toward the bottom income quintile across all four countries. Hence, if QE leads to a considerable reduction of aggregate unemployment, a proportionally larger number of individuals in the lower income quintiles are drawn out of unemployment and this tends to reduce income inequality. We find that this second effect dominates and, hence, the stimulative effects of QE on employment are skewed toward low-income households. Across the four countries, the peak unemployment response for the bottom income quintile ranges roughly between $-1\%$ and $-2\%$, while for the highest income quintile unemployment declines by less than $0.5\%$.

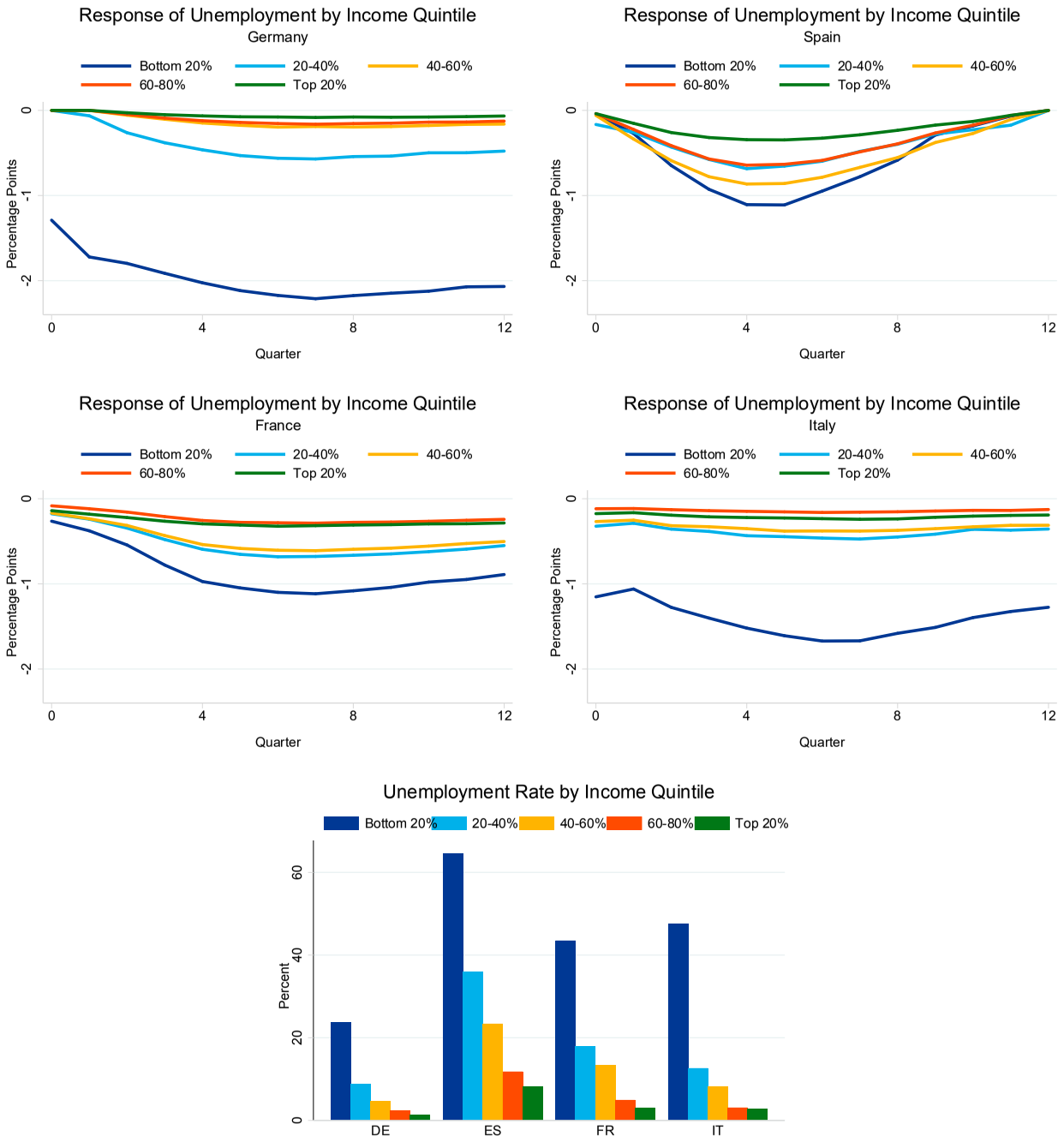
The micro impulse responses also vary across countries, both regarding the level and the dispersion of responses across income quintiles (Figure 4). One factor to explain the differences, in particular for the levels, is the cross-country difference in macro responses. For example, the overall reduction in unemployment is larger in Spain than in the other three countries. Instead, the dispersion of micro impulse responses across income quintiles is importantly affected by the distribution of the unemployed across income quintiles, which varies across countries. Indeed, a relevant mass of unemployed people in Spain lives in households whose income falls into higher quintiles, so that the differences in impulse responses across quintiles in Spain are smaller (see, again, the bottom right panel in Figure 4). In contrast, the number of the unemployed in Germany and Italy is more strongly skewed toward the lowest income quintile, which causes unemployment in the lowest income quintile to drop more (relative to other quintiles).[19]

Figure 5 shows the micro responses of mean income by income quintile, combining the earnings heterogeneity and the income composition channels. These responses are primarily driven by the transitions into employment and by differences in unemployment

[18] In order to appreciate the quantitative relevance of this heterogeneity in probabilities to become employed, a counterfactual scenario where all individuals have the same probability to be drawn out of unemployment implies a significantly stronger stimulating effects on the lower income quintiles compared to our scenario based on estimated probabilities—as documented in the online Appendix C.

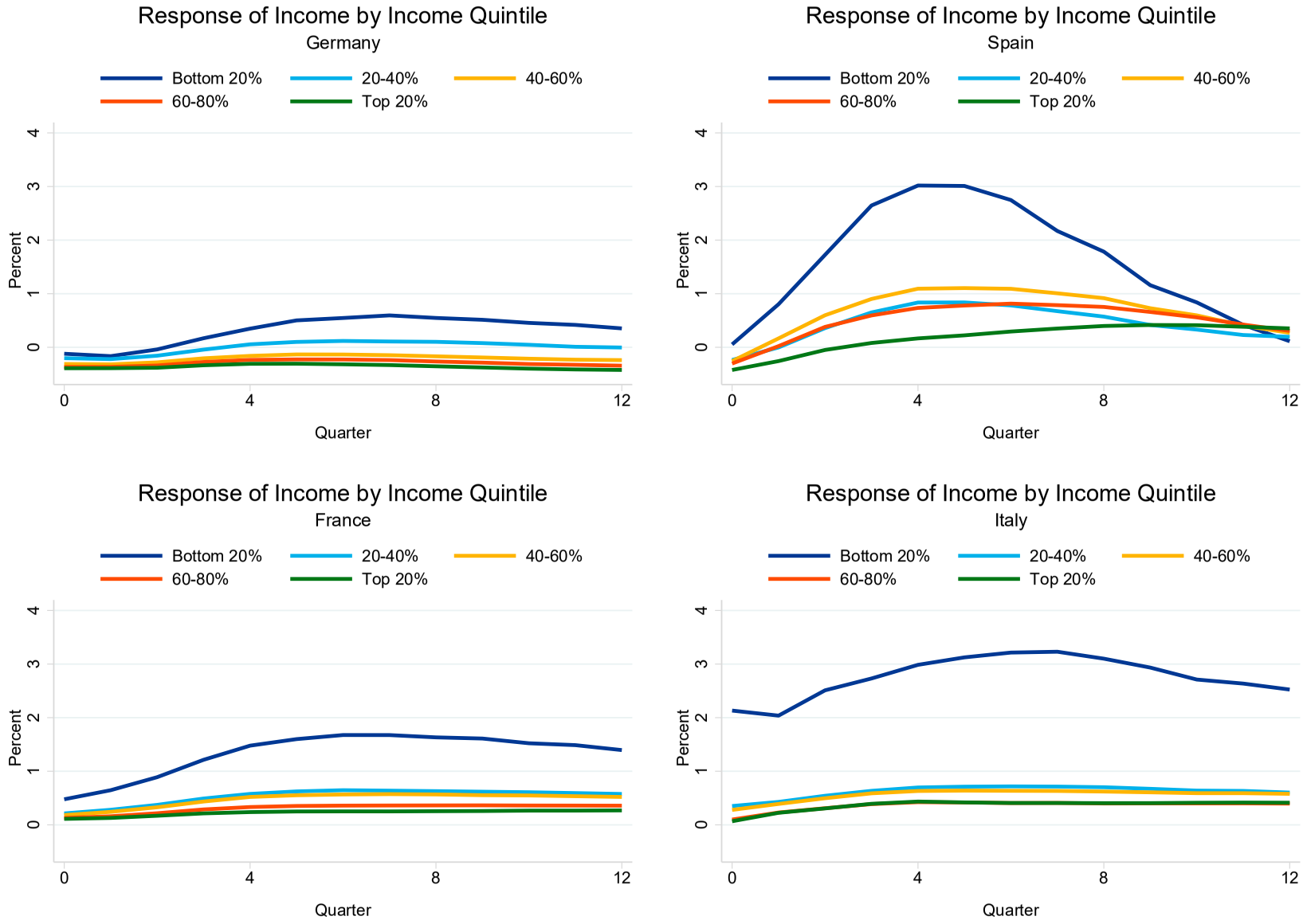
[19] Notice also that, in principle, the cross-country dispersion could also be explained by the fact that the employment probabilities in the probit models (1) are country-specific. This factor turns out to play a very minor role to explain our results.

**Figure 4** Impulse Responses of Unemployment by Country and Income Quintile




**Source**: Household Finance and Consumption Survey, wave 2014
**Note**: The figures show impulse responses of unemployment by income quintile.

**Figure 5** Impulse Responses of Mean Income by Country and Income Quintile




**Source**: Household Finance and Consumption Survey, wave 2014
**Note**: The charts show impulse responses of mean income by income quintile.

**Table 2** Effects of Quantitative Easing on Income and Wealth Inequality

| | Gini Coefficient (%) | |
|---|---|---|
| | Income | Net Wealth |
| Actual Data | 43.161 | 69.168 |
| Baseline Simulation | 43.034 | 69.177 |
| Robustness Scenarios | | |
| 1. Effects of Financial Income (2.1%) | 43.059 | |
| 2. Effects of Financial Income (Country-Specific Response 3 to 8%) | 43.066 | |
| 3. Stock Trading (Household- & Country-Specific Response 2 to 7%) | | 69.204 |
| 4. Local House Prices (Country-Specific Response −1 to 3%) | | 69.225 |

The table shows the Gini coefficients for gross household income and net wealth for actual data, the baseline scenario and four alternative, robustness scenarios described in section 3.2.3: two scenarios accounting for the effects of financial income, a scenario on portfolio rebalancing of stocks (stock trading) and a scenario with heterogeneity in responses of house price to quantitative easing. The scenarios report the Gini coefficients four quarters after the impact of the quantitative easing shock.

insurance replacement rates (as estimated by the Heckman model). The replacement rates are in general more generous in Germany and France than in Spain and, in particular, Italy.[20] As a result, the magnitude and dispersion of income responses in Italy and Spain is larger. For example, the large positive response in mean income of the lowest quintile in Italy arises thanks to both the substantial decline in unemployment rate highlighted in Figure 4 and the substantial increase in (labor) income of the newly employed individuals.[21]

These findings imply that the earnings heterogeneity channel is the most relevant factor to explain the changes in income across quintiles. To more precisely show this point, Figure 6 decomposes the overall increase in mean income into the extensive (earnings heterogeneity) and the intensive margins (income composition) for an aggregate of the four countries, one year after the shock. The extensive margin is particularly strong in the bottom income quintile, for which wage growth plays a very small role. However, transitions from unemployment to employment account for the bulk of the total increase on income across much of the whole distribution (except for the top income quintile). The role of the intensive margin is small due to the relatively muted response of wages to the monetary policy shock.[22]

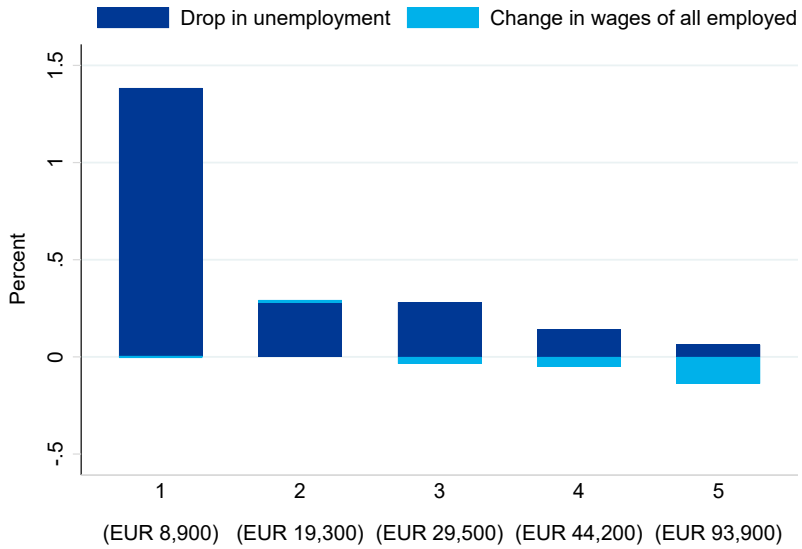
To summarize the effects on income inequality, Table 2 shows that quantitative easing reduces the Gini coefficient for gross household income from 43.16%, the value computed

[20] See, e.g., the OECD Statistics on Benefits and Wages: http://www.oecd.org/els/benefits-and-wages-statistics.htm.

[21] The results are shown for gross (pre-tax) income. The increase in after-tax income would be somewhat lower, however, not by much, as most newly employed people are not subject to large taxes. As for the effect on inequality of net income, it would be reduced more than inequality of gross income because of progressivity of taxes.

[22] For two countries the response of aggregate wages is negative, which translates into a negative impact of the intensive margin in the upper quintiles of Figure 6.

**Figure 6** Decomposition of the Total Effect on Mean Income into the Extensive and the Intensive Margin




**Source**: Household Finance and Consumption Survey, wave 2014
**Note**: The figure shows the percentage change in mean income across income quintiles in the euro area four quarters after the impact of the QE shock. It also shows the decomposition of the change into the extensive margin (transition from unemployment to employment) and the intensive margin (change in wage). The numbers in parentheses show the initial levels of mean gross household income. The figure shows an aggregate of France, Germany, Italy and Spain.

from the HFCS in 2014, to 43.03%. These results are in line with recent analysis on the sensitivity of individual incomes to business cycle and monetary policy. For example, Alves et al. (2019) and Heathcote et al. (2020) estimate that individuals with lower earnings are particularly sensitive to aggregate fluctuations. Similarly, Broer et al. (2020) find in German data that earnings in the bottom tail of the distribution are particularly sensitive to monetary policy shocks. In addition, they report that income risk for the poor is almost entirely extensive (due to labor market transitions), while at higher incomes intensive risk is much more important.[23] These empirical results are also consistent with models with indivisible labor, which imply higher labor supply elasticities for lower income groups (e.g., Chang and Kim, 2007 and Ma, 2020).

### *3.2.2 Effects on Household Wealth: The Portfolio Composition Channel*

This section analyses how the portfolio composition channel affects household net wealth. Figure 7 shows the micro responses of mean net wealth by wealth quintile. These responses arise from a combination of the response of house prices, stock prices and bond prices, and holdings of wealth components across the distribution (and countries).

Broadly, the responses of wealth in quintiles two to five increase by around 1.5% in France, Spain and Italy, and are rather flat in Germany. There is little evidence that the median wealth among the top wealth quintile households would increase more strongly, though this does happen for the top 10% of the wealth distribution, where the holdings of stocks are prevalent. Overall, Table 2 documents that the Gini coefficient for net wealth is only modestly affected by QE, rising from 69.17% to 69.18%. An important takeaway from this exercise is the key role of including house prices in the analysis, since most households own large holdings of housing wealth rather than stocks and bonds, which are only relatively more prominent in the top tail of the distribution.[24]

### *3.2.3 Robustness Checks*

This section explores whether some plausible perturbations of our baseline specification affect the main results. In particular, we extract the time series of the QE shocks from our multi-country VAR and we use the local linear projection method of Jordà (2005) to derive the effects of the shock on additional variables which are useful to capture alternative scenarios to our baseline analysis.[25]

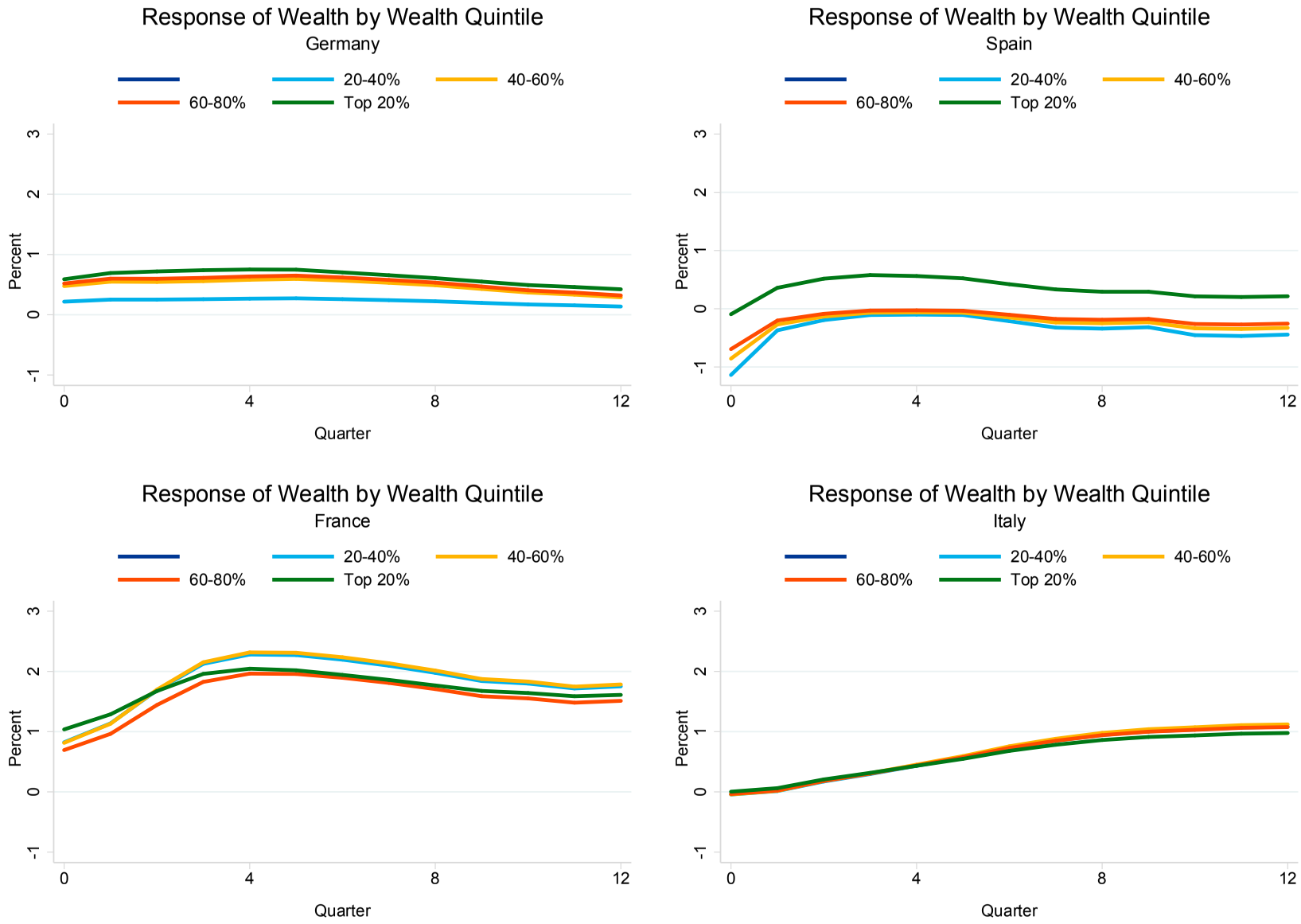
Our baseline analysis of the income composition channel neglects the effects of QE on financial income, which is disproportionately earned by the top tail of the income

[23] For the US, Guvenen et al. (2017) estimate a U-shaped exposure of individual earnings to (aggregate) GDP growth, which is rising in the top tail, above the 99th percentile of earnings. Unfortunately, to our knowledge, corresponding evidence on the exposure for the highest percentiles from European countries is not available because of data limitations. In section 3.2.3 we consider a parametrization for which monetary policy stimulates financial income in line with our estimates from aggregate data.

[24] This finding is in line with Adam and Tzamourani (2016); see, e.g., their Figure 4. See also Kuhn et al. (2020), Figure 17 for historical evidence from the US and Martínez-Toledano (2020) for estimates from Spain.

[25] See Appendices A and B for the description of the local linear projection method and for more information on the data sources we use for the robustness checks.

**Figure 7** Impulse Responses of Mean Net Wealth by Country and Net Wealth Quintile




**Source**: Household Finance and Consumption Survey, wave 2014
**Note**: The figures show impulse responses of net wealth. The response for the bottom 20% not shown as the value of net wealth in the lowest quintile is negative or close to EUR 0.

distribution.[26] If quantitative easing increases financial income, e.g., via stimulating corporate profits, this effect may widen income inequality.[27]

To address this issue, we estimate the responses to the QE shock of two alternative (aggregate) measures of financial income: (i) profits (available on the euro area level) and (ii) net property income (available for the four countries under analysis).[28] The top panels of Figure 8 consider the implications for the income distribution of two scenarios based on such responses: (i) assuming that financial income behaves similarly to profits (i.e., increasing by about 2% in all countries; left panel), and (ii) assuming that financial income responds as estimated by linear projections for aggregate data on net property income (right panel). As expected, the scenarios increase income in particular among the top income quintile of households, but only by an additional 0.2% or less. As a result, the overall impact on total income is quite limited, changing the Gini coefficients for the two scenarios only marginally compared to our baseline assessment, falling to roughly 43.06% (instead of 43.03% estimated for the baseline; see Table 2, robustness scenarios 1 and 2).

Turning to wealth inequality, we first relax our assumption of no portfolio rebalancing. To get an idea of a plausible amount of rebalancing, we rely on aggregate country-level flow-of-funds data on the holdings of different asset categories by households, and we compute their response to a QE shock by means of local projections. Then, in line with the responses of these variables to the shock, we simulate a scenario in which quantitative easing increases aggregate holdings of stocks by 2% to 7% across the four countries. In addition, in line with the household finance literature on trading,[29] we account for more reallocation by wealthy households, which are likely to be more attentive and trade more, by calibrating that their stock holdings to increase by 2% more than the holdings of households in the middle quintile.

We find that stock trading affects the distribution of net wealth only very little (Figure 8, bottom left panel), increasing mean wealth in the top quintile by about 0.2%. Correspondingly, Table 2 documents that the Gini coefficient on wealth under this alternative scenario (robustness scenario 3) rises to 69.20% (compared to 68.18% for the baseline scenario). This is explained by the fact that the share of stocks in the portfolios of European households lies below 5% even for the top wealth quintile. We view this finding as an upper bound of the extent to which active portfolio rebalancing can affect wealth inequality because evidence from micro data, including the influential work of Calvet et al. (2009) and Brunnermeier and Nagel (2008), typically estimates that (if at all) individual households tend to actively rebalance in the opposite direction, i.e., by selling (not buying) risky financial assets after experiencing high returns.

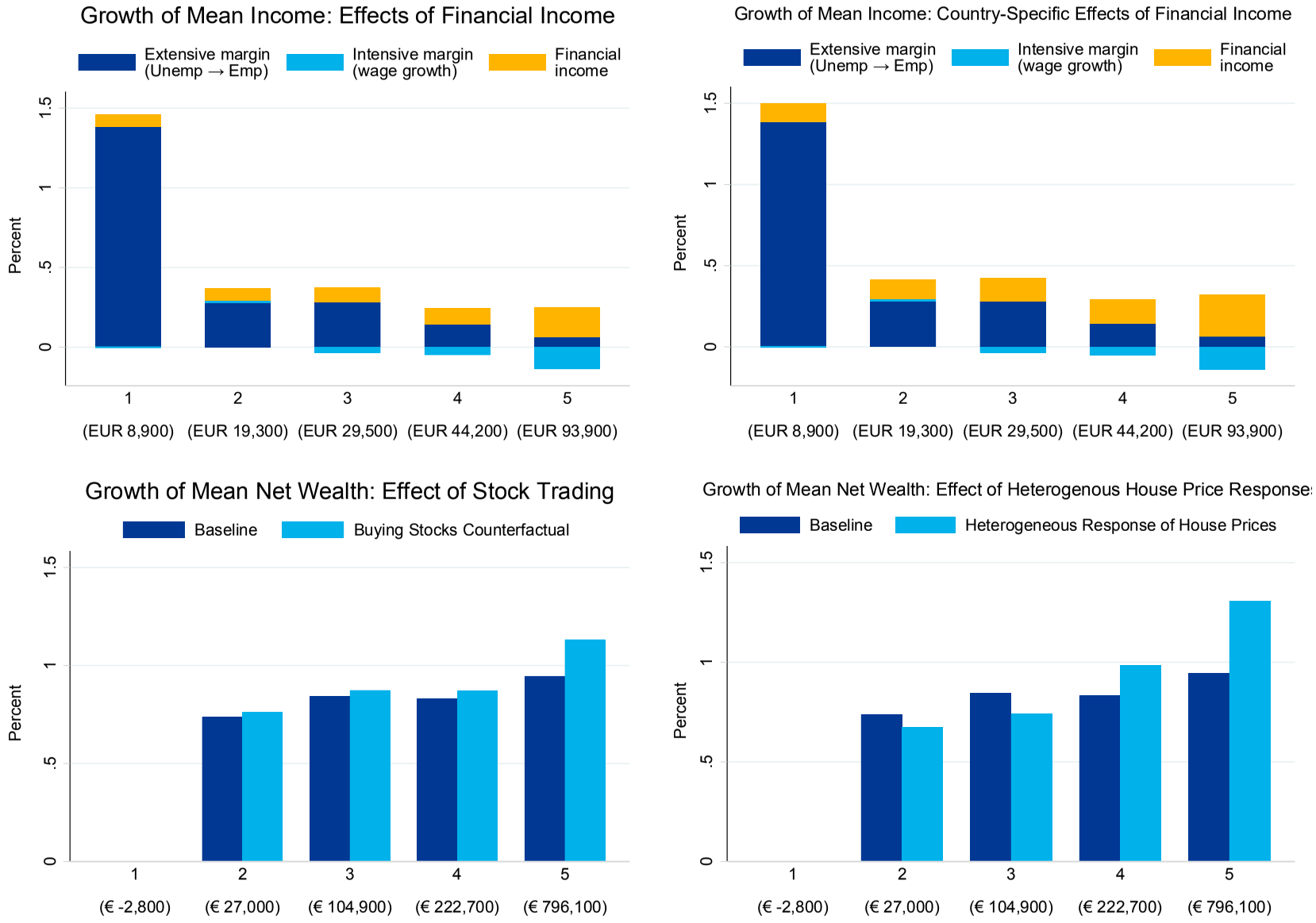
[26] Financial income includes income in the form of interest or dividends on sight deposits, time and saving deposits, certificates of deposit, managed accounts, bonds, publicly traded stock shares or mutual funds. More broadly, we also include income from renting real estate and income from private business other than self-employment.

[27] Existing evidence, e.g., Guvenen et al. (2014), points to slight, rather than strong, pro-cyclicality in the unconditional dynamics of earnings and financial income among top earners.

[28] See the online Appendix C for the results.

[29] For example, Calvet et al. (2009) find that more educated and wealthier households tend to rebalance their portfolios more actively. Similar, Bilias et al. (2010) report that households with higher education, income and net financial wealth trade more.

**Figure 8** Robustness Scenarios for Income and Wealth




**Note**: The figures show for various scenarios the percentage increase in mean income and mean net wealth across quintiles of gross household income and net wealth for an aggregate of France, Germany, Italy and Spain.
The top left panel shows the implications for gross income of an increase of 2.1% in financial income. The top right panel shows the implications for gross income of country-specific increase in financial income (France: 3.7%, Germany: 3.5%, Italy: 3.6%, Spain: 8.0%; bottom panel).
The bottom left panel compares the baseline scenario with the one in which the holding of stocks increases for the middle net wealth quintile in each country as follows: France 6.2%, Germany 3.2%, Italy 2.0%, Spain 7.3%. In addition, across wealth quintiles the spread varies by ±2%, so that households in the lowest quintile buy 2% less and households in the highest quintile buy 2% more stocks than those in the middle quintile.
The bottom right panel compares the baseline scenario with the one in which house prices of more expensive houses (in terms of price per square meter) react more strongly to monetary policy. House price growth ranges across quintiles of price per square meter as follows: Germany, Italy and Spain −0.5% to 1.5% and France 1.0% to 3.0%.
The numbers in parentheses show the initial levels of mean net wealth. The change for the lowest quintile of net wealth is not shown because its initial level is negative.

Finally, our treatment of wealth inequality does not account for a possible heterogeneity in the responses of house prices across regions (arising, e.g., due to differences in elasticity of housing supply). For this reason, we also investigate a scenario in which the prices of more expensive houses (measured in EUR per square meter) react more strongly to quantitative easing. This calibration is based on our estimates exploiting regional data from Spain, which suggest that the prices of more expensive houses respond more strongly to monetary policy.[30] Specifically, the dispersion in the responses of house prices across Spanish provinces is about 3 percentage points. Figure 8 shows the comparison of our baseline results with those obtained by assuming that the increase in house prices due to the QE shock also depends on the level of house prices.[31] Because poorer households tend to own less expensive houses, this alternative assumption increases the dispersion of growth rates of net wealth across quintiles: for the second lowest net wealth quintile, mean wealth grows by 0.7% (slightly less than for the baseline), while for the top quintile the mean wealth increases by about 1.3% (compared to around 1% for the baseline). Table 2 shows that under this scenario the Gini coefficient for net wealth rises from 69.17% to 69.23% (compared to 69.18% for the baseline) and, hence, our conclusion that the effect of quantitative easing on wealth inequality is overall quite muted remains unaffected.

# 4 Conclusions

Combining estimates from a VAR with aggregate data and a simulation on household-level data, we quantify how quantitative easing in the euro area affect individual households via the portfolio composition, the income composition and the earnings heterogeneity channels. We find that although QE has only negligible effects on wealth inequality, it noticeably compresses the income distribution since many households with lower incomes become employed. Specifically, a year after the shock, the Gini coefficient for income falls from 43.16% to 43.03%, while the change of the Gini coefficient for net wealth is an order of magnitude smaller.

The effects of monetary policy fade away over time and, hence, quantitative easing should not be a key driver of inequality in the long run, when other factors, such as globalization or the progressivity of the tax system are more important. However, our results suggest that quantitative easing substantially contributed to support vulnerable households.

Our results are also informative about the strength and nature of the transmission of monetary policy to consumption. An extensive literature has recently documented that constrained households—e.g., those with low incomes or little liquid assets—have high marginal propensities to consume. We find such households also particularly benefit

[30] See Figure C.3 in the online Appendix C. Spain is the only country in our sample for which quarterly data on regional house prices are available since 1999. Fagereng et al. (2020) estimates positive *unconditional* correlation between the level of wealth and returns to wealth in Norwegian data.

[31] Specifically, house price growth ranges across quintiles of price per square meter as follows: Germany, Italy and Spain −0.5% to 1.5% and France 1.0% to 3.0%. This calibration thus preserves the aggregate response of house prices to quantitative easing estimated in the VAR, upper right-hand panel in Figure 3, and adds to it a positive relationship between the level of house prices and their sensitivity to monetary policy.

from a monetary stimulus, which boosts their employment and income. In combination, these two facts imply that the stimulating effect of quantitative easing on aggregate consumption is substantially amplified both because it disproportionately boosts incomes in the lower part of the distribution and because this impulse has a stronger effect on consumption via the larger MPCs of the constrained households.[32]

[32] Patterson (2019) documents a positive covariance between worker MPCs and the elasticity of their earnings to GDP in the US data. Slacalek et al. (2020) quantify the channels of monetary transmission to consumption and their heterogeneity across households.

# Appendix A: Estimation

## A.1 The Prior Distributions

The prior distributions in our Bayesian VAR are specified as follows. For the prior on the covariance matrix of the errors, we set the degrees of freedom of the Inverse Wishart distribution equal to $N + 2$ (where $N = 22$ is the number of variables in the model), the minimum value that guarantees the existence of the prior mean, and we assume a diagonal scaling matrix $\Psi$, which we treat as a hyperparameter.

The baseline prior on the model coefficients is a version of the Minnesota prior (see Litterman, 1979). This prior is centered on the assumption that each variable follows an independent random walk process, possibly with drift. The prior first and second moments for the VAR coefficients are:

$$\begin{aligned}
\mathbf{E}\Big((B_s)_{ij}\Big|\Sigma\Big) &= \begin{cases} 1 & \text{if } i=j \text{ and } s=1 \\ 0 & \text{otherwise} \end{cases}, \\
\operatorname{cov}\Big((B_s)_{ij},(B_r)_{hm}\Big|\Sigma\Big) &= \begin{cases} \lambda^2 \frac{1}{s^2} \frac{\Sigma_{ih}}{\psi_j/(d-n-1)} & \text{if } m=j \text{ and } r=s \\ 0 & \text{otherwise} \end{cases}.
\end{aligned}$$

Notice that the variance of this prior is lower for the coefficients associated with more distant lags and that coefficients associated with the same variable and lag in different equations are allowed to be correlated. Finally, the key hyperparameter is $\lambda$, which controls the scale of all variances and covariances and effectively determines the overall tightness of this prior. The terms $\Sigma_{ih}/\Psi_j$ account for the relative scale of the variables. The prior for the intercept $C$ is non-informative.

The Minnesota prior is complemented by a prior on the sum of the VAR coefficients, introduced as refinements of the Minnesota prior to further "favor unit roots and cointegration, which fits the beliefs reflected in the practices of many applied macroeconomists" (see Sims and Zha, 1998, p. 958). This additional prior tends to reduce the importance of the deterministic component implied by VARs estimated conditioning on the initial observations (see Sims, 1996 and Giannone et al., 2019). This prior is known as no-cointegration (or, simply, sum-of-coefficients) prior.

To understand what this prior entails, we rewrite the VAR equation in an error-correction form:

$$\Delta y_t = C + (B_1 + \cdots + B_p - I_N) y_{t-p} + A_1 \Delta y_{t-1} + \cdots + A_p \Delta y_{t-p} + \epsilon_t,$$

where $A_s = -B_{s+1} - \cdots - B_p$. A VAR in first differences implies the restriction $\Pi = (B_1 + \cdots + B_p - I_N) = 0$. Doan et al. (1984) introduced the no-cointegration prior which centered at 1 the sum of coefficients on own lags for each variable, and at 0 the sum of coefficients on other variables' lags. This prior also introduces correlation among the coefficients on each variable in each equation. The tightness of this additional prior is controlled by the hyperparameter $\mu$. As $\mu$ goes to infinity, the prior becomes diffuse, while as it goes to 0, it implies the presence of a unit root in each equation.

The setting of the prior distributions depends on the hyperparameters $\lambda$, $\mu$ and $\Psi$, which describe the informativeness of the prior distributions for the model coefficients.

In setting these parameters, we follow the theoretically grounded approach proposed by Giannone et al. (2015), who suggest to treat the hyperparameters as additional parameters, in the spirit of hierarchical modelling. As hyper-priors (i.e., prior distributions for the hyperparameters), we use proper but almost flat distributions.

### A.2 The Local Linear Projection

Our robustness exercises in section 3.2.3 adopt the local linear projection to derive the response of various variables to the shocks we estimate in the VAR. Let us briefly describe our application of the method developed in Jordà (2005). Denote $G_t$ an additional variable of interest. We transform these variables as for the VAR, i.e., we compute annualized log-levels unless the variable is already expressed in terms of rates. Denote as $g_t$ the transformed variable.

We evaluate the impulse response $\vartheta^h$ of $g_t$ to the shock $\varepsilon_{1,t}$ at the horizon $h$ by regressing $g_{t+h}$ on $\varepsilon_{1,t}$ and the lags of $g_t$. Specifically, we estimate the following regression:

$$g_{t+h} = \alpha + \vartheta^h \varepsilon_{1,t} + \gamma(L) g_t + \eta_t.$$

The regression is estimated by means of Bayesian techniques. We impose a flat prior on $\alpha$ and $\vartheta^h$, while we impose an informative prior on the coefficients on the lags, $\gamma(L)$. The informative prior has the exact same features of the Minnesota prior described in Appendix A. Notably, the shrinkage of the lagged terms grows with the horizon $h$ at which the impulse response is computed.

## Appendix B: Macroeconomic Data

**Table 3** Macroeconomic Database

| **Variable** | **Transformation** | **Source** |
|---|---|---|
| **France, Germany, Italy, Spain** | | |
| GDP | log-levels | Eurostat |
| GDP deflator | log-levels | Eurostat |
| Unemployment rate | levels | Eurostat |
| House prices | log-levels | Eurostat |
| Compensation per employee | log-levels | Eurostat |
| **Euro area** | | |
| Long-term interest rate | levels | AWM database (LTN) |
| Stock prices | log-levels | ECB SDW |

Table 3 describes our aggregate time series.

In our robustness exercises, we exploit some additional data sources, available at quarterly frequency for the sample 1999Q1–2016Q4. First, we consider data on profits

for the euro area; precisely, this variable captures gross operating surplus (total economy, nominal, seasonally adjusted data) and is available from the Main National Accounts collection in the ECB Statistical Data Warehouse (SDW). The data on net property income and stock holdings of the four countries under analysis come from the Euro Area Sectoral Accounts. Finally, the data on regional house prices in Spain are available from the website of the Spanish government, Ministerio de Fomento.[33]

[33] We use the series "valor tasado medio de vivienda libre" (the aggregate house price, total national, and the house prices of the 17 regions for which the quarterly data are available, i.e., we exclude the autonomous cities Ceuta and Melilla): http://www.fomento.gob.es/BE2/?nivel=2&orden=35000000.

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