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**Conference Paper**

## Public Insurance and Wealth Inequality - A Euro Area Analysis

Beiträge zur Jahrestagung des Vereins für Socialpolitik 2016: Demographischer Wandel - Session: Inequality in Germany and Europe, No. A06-V3

**Provided in Cooperation with:**

Verein für Socialpolitik / German Economic Association

*Suggested Citation:* Pham-Dao, Lien (2016) : Public Insurance and Wealth Inequality - A Euro Area Analysis, Beiträge zur Jahrestagung des Vereins für Socialpolitik 2016: Demographischer Wandel - Session: Inequality in Germany and Europe, No. A06-V3, ZBW - Deutsche Zentralbibliothek für Wirtschaftswissenschaften, Leibniz-Informationszentrum Wirtschaft, Kiel und Hamburg

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# Public Insurance and Wealth Inequality: A Euro Area Analysis\*

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July 4, 2016

## Abstract

This paper assesses the quantitative importance of cross-country differences in labor market dynamics and social security institutions for euro area differences in private net wealth inequality. I document the empirical puzzle that euro area countries with the largest reduction in the income Gini coefficient through public transfers robustly show higher inequality in private net wealth. Revisiting the argument by Hubbard et al. (1995) that public insurance crowds out private savings especially of the poor, I construct a life-cycle model with heterogeneous households and incomplete markets that features exogenous labor market risks, unemployment benefits, mean-tested minimum income support and public and occupational pensions. Calibrating the model to the euro area differences in the net earnings process, unemployment dynamics and social security system, it can account for 70.1% of the cross-country differences in the net wealth Gini coefficients for the bottom 95% of the wealth distribution. Welfare policies contribute 57.5% to the wealth inequality differences across the euro area, while net earnings and unemployment dynamics account for 12.6%.

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\*I would like to thank Christian Bayer, Anmol Bhandari, Fatih Guvenen, Mariacristina De Nardi, Thomas Hintermaier, Tim Kehoe, Fabian Kindermann, Moritz Kuhn, Keith Kuester, Alexander Ludwig, Ellen McGrattan, Felix Wellschmied and the Household Finance and Consumption Network (HFCN) for their helpful comments. In particular, I thank Felix Wellschmied for sharing his code. The research leading to these results has received funding from the European Research Council under the European Union's Seventh Framework Programme (FTP/2007-2013) / ERC Grant agreement no. 282740. This paper uses data from the HFCS and EU-SILC. The results published and the related observations and conclusions are mine and do not correspond to results or analysis of the HFCN, Eurostat, the European Commission or any other national statistical authorities.

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**Keywords:** Wealth inequality, public insurance, transfers, cross-country analysis.  
**JEL-Codes:** D31, D91, E21, H31

# 1 Introduction

The first wave of the Household Finance and Consumption Survey (HFCS), mainly conducted in the period from 2009 to 2011, reveals that there are large cross-country variations in household private net wealth inequality for the ten largest economies in the euro area. The Gini coefficient of household net wealth ranges from 0.76 for Austria to 0.56 in Greece. Since the release of the HFCS in 2013, the causes of the large euro area differences in private net wealth inequality and the surprisingly low median wealth in some euro area countries with high GDP per capita have been at the forefront of public and political debates. In particular, there has been a discussion about the role played by institutional factors. This paper aims to contribute to this debate by assessing the quantitative importance of cross-country differences in social security institutions and labor market dynamics for euro area differences in private net wealth inequality, as measured by the Gini coefficient.

The interest in social security as a potential determinant of the wealth distribution is motivated by the surprising finding that those countries with a larger reduction in the income Gini coefficient through redistributive public transfers to households robustly show higher inequality in private net wealth. On its own, this correlation falls short of providing a quantitative assessment of the importance of these policies. Therefore, going back to the theoretical contribution by Hubbard et al. (1995) that public insurance crowds out private savings, especially of the poor, I construct a life cycle model with heterogeneous households and incomplete markets that features exogenous labor market risks, unemployment insurance, means-tested minimum income support, and pension benefits. There are three key determinants for wealth accumulation in the model, namely old-age provision, leaving bequests, and precautionary savings to self-insure against gross earnings, unemployment and life-span risk. The more redistributive transfers are across households, the more pronounced the crowding-out effect is on private precautionary savings for low-income households, as there is a relatively greater reduction in their need for self-insurance.

Calibrating the model to the euro area differences in net earnings processes and unemployment dynamics and to the institutional differences in public insurance, the model can account for 70.1% of the euro area variation in private net wealth inequality. Moreover, by adopting a modified method based on Guvenen et al. (2014), I provide a decomposition that disentangles how much each factor contributes individually to this fraction. The model results suggest that welfare policies contribute 57.5% to euro area differences in net wealth Gini coefficients for the bottom 95% of the wealth distribution. It turns out

that the most important institution of the social security system for determining wealth inequality differences across the euro area is means-tested minimum income support. It contributes the lion's share of 44.8%, followed by public and occupational pension schemes, which account for 10.7%. Differences in unemployment benefits, by contrast, play only a minor role, contributing 2%. Furthermore, the net earnings process and unemployment dynamics can jointly rationalize 12.6% of the euro area differences in private net wealth inequality. This important role of public insurance for wealth inequality patterns in the euro area also sheds light on why welfare states with more redistributive transfers show higher wealth inequality. While transfers directly mitigate income differences across households, their general availability leads to a more unequal wealth distribution in the long run.

The importance of minimum income support programs for the wealth distribution relative to other policies is rooted in several distinct features. First, minimum income benefits are not dependent on past contributions and hence are more redistributive across individuals compared to unemployment benefits or pensions, which are instead more redistributive over the life cycle in the euro area. The lower bound on consumption leaves households with high expected life-time income relatively unaffected in their precautionary savings decision, while the need for self-insurance of households in the lower part of the income distribution is substantially reduced, thereby increasing wealth inequality. Second, minimum income assistance guarantees a certain lump sum transfer, while future potential unemployment benefits replace a constant fraction of previous net income and are hence still dependent on uncertain net income. Similarly, there remains some uncertainty about the exact pension level during retirement, as it will depend on the household's pre-retirement labor market performance. As households are risk averse, the effects of unemployment and pension benefits on wealth inequality turn out weaker despite strong crowding out effects on aggregate private savings. Third, the asset-based means-testing of minimum income support introduces an implicit tax on savings, such that low-wealth households face a trade-off between saving for bad income states and dissaving to become eligible for income support. And last, minimum income benefits are of unlimited duration, while unemployment benefits mitigate earnings losses only temporarily.

Regarding the calibration of the model, I assume that all countries share the same key technological and preference parameter values and allow them to differ only in the parameters describing the unemployment and net earnings process, as well as the social security system. The coefficient of determination then provides a measure to quantify what fraction of cross-country differences in wealth inequality are generated solely by

these country-specific features of the labor market process and social security institutions. For the decomposition exercise, I construct a fictive euro area country as a reference unit, whose parameters correspond to the average of the individual countries' parameters. I then sequentially set the country-specific parameters to the parameters of the constructed reference euro area country until no parameter differences between countries remain, and each time determine the explanatory power of the model. This way, I can quantify how much each factor contributes individually to the overall fraction of cross-country differences explained by the model. The variances of the gross earnings processes are estimated from household income data of the European Survey on Income and Living Conditions (EU-SILC) for 2004 until 2010. The calibration of labor income tax schedules and welfare policies is mainly based on estimates from the OECD benefit and tax model, or own estimates from the EU-SILC.

While several studies have highlighted the distortionary effects of certain institutions of the social security system on aggregate savings, less research has concentrated on the consequences of public insurance for private net wealth inequality. This is because research on wealth inequality has so far mainly focused on the upper tail of the wealth distribution. In contrast, this paper sheds light on the remaining part of the wealth distribution. It is shown that, when considering the bottom 95% of the wealth distribution, large euro area differences in wealth inequality remain, and that public transfers and labor market dynamics are indeed central for determining wealth inequality patterns across the euro area.

The paper is organized as follows. Section 2 discusses the current literature, and section 3 presents empirical evidence. Section 4 introduces the model, section 5 the calibration strategy, and section 6 presents the quantitative results. Section 7 concludes.

## 2 Related Literature

There is a large theoretical literature on the determinants of wealth inequality, with a particular focus on the high wealth concentration in the United States. In fact, the vast majority of the work has focused on the upper tail of the wealth distribution and, in particular, on explaining high savings rates of wealthy people documented in the empirical literature (Dynan et al., 2000). To account for these, standard incomplete markets models have been extended by various features such as heterogeneity in patience (Krusell and Smith, 1998), transmission in human capital and voluntary bequests across generations (De Nardi, 2004), entrepreneurship (Quadrini et al., 1999), high returns on capital in the presence of borrowing constraints (Cagetti and De Nardi, 2006) or high

earnings risk for top earners (Castaneda et al., 2003).

In contrast, less research has concentrated on the remaining part of the wealth distribution and, in particular, on the low savings of income poor households. In a seminal paper, Hubbard et al. (1995) highlight the distortionary effects of social insurance on households savings behavior through the reduction of income risk and an implicit tax on savings in the presence of asset means-testing. In particular, they aim to explain why many low-income households accumulate only little wealth over the life cycle, much less than a standard life cycle model would suggest. Other work relates precautionary savings to other institutions of social security systems, such as health insurance (Kotlikoff, 1986), public pension systems or unemployment insurance. Engen and Gruber (2001) show theoretically and empirically that higher unemployment replacement rates crowd out aggregate savings. For the empirical analysis, they exploit differences in unemployment generosity across U.S. states.

While most of those papers shed light on the distortion of aggregate savings through specific institutions of the social security system, I jointly model three of them and analyze their implications for wealth inequality. In a similar vein, two papers explicitly relate wealth inequality to public insurance and more concretely public pension systems for specific countries. Domeij and Klein (2002) show that wealth inequality in Sweden is driven to a large extent by its very redistributive public pension scheme. Given the current discussion about wealth inequality fueled by Thomas Piketty, Kaymak and Poschke (2016) made a recent contribution, analyzing the extent to which institutional changes from 1960 to 2010 in the U.S. can explain the increased share of wealth held by the top percentiles. They find that despite the dominant role of changing wage inequality, the expansion of social security in terms of more generous pensions and Medicare can account for an important portion. Hintermaier and Koeniger (2011) also attribute changes in the U.S. net wealth distribution across time to increases in income risk. Considering similar savings motives in an incomplete markets life-cycle model as in the present paper, they also abstract from the upper tail of the wealth distribution for their analysis.

Regarding the euro area, Fessler and Schuerz (2015) provide empirical evidence based on the HFCS for the role of welfare state policies in explaining cross-country differences in household net wealth. Controlling for various household characteristics and inheritance, they find in a multilevel cross-country regression that welfare state expenditures across countries are negatively correlated with household net wealth and hence a substitute for private wealth. Moreover, they show that the substitution effect of pension and social security expenditures with regard to private wealth holdings is significant along the wealth distribution, but relatively lower at high wealth levels. While the latter result

empirically confirms the hypothesis that more generous public insurance increases wealth inequality, this present paper explicitly models various features of the social security systems in the euro area and quantifies and decomposes their importance for cross-country differences in wealth inequality.

The only other paper explicitly analyzing wealth inequality differences in the euro area is empirical in nature and has stressed the importance of cross-country differences in home ownership rates (Kaas et al., 2015) and the fact that tenants accumulate less wealth on average than homeowners. I see my paper as complementary to their work, as investment in housing is one way to accumulate wealth. Furthermore, given that empirical evidence suggests a positive correlation of household income and home ownership status and that tenants are commonly low-income households, social security might shed light on the still puzzling question of why tenants accumulate so little wealth compared to homeowners.

### **3 Redistributive Policies and Wealth Inequality in the Euro Area**

#### **3.1 Cross-country Evidence on Wealth Inequality and the Degree of Redistribution of Public Transfers**

The newly available household data on private wealth from the Household Finance and Consumption Survey conducted by the ECB (2013) allows a euro area wide comparison of household wealth, given the ex ante coordination of the survey questionnaire and methodology and its emphasis on output harmonization. As in the Survey of Consumer Finances, oversampling procedures of wealthy households are applied in order to achieve unbiased estimates of wealth and its distribution. The release of the first wave of the HFCS in 2013, covering household interviews mainly conducted in the period from 2009 to 2011, allows a reasonable comparative analysis of the distribution of household wealth across the euro area.<sup>1</sup>

First, I document along the vertical axis of Figure 1 that there are large cross-country variations in household private net wealth inequality for the ten largest economies in the euro area.<sup>2</sup> Household net wealth is defined as the household's total assets, i.e.

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<sup>1</sup>The data for Spain in the first wave refers to the year 2008. Since the Spanish survey of household finances (EFF) is the only survey which was conducted before the financial crisis, I use already published data from the second wave which corresponds to the year 2011 and hence allows for a post-crisis analysis in all countries. A comparison of the wealth Gini coefficients for the two waves in Spain indicates an increase in wealth inequality from 0.581 to 0.608 after the bursting of the housing bubble.

<sup>2</sup>In my analysis, I focus on ten euro area countries, for which I have complete data on all parameters of interest: Austria (AT), Belgium (BE), Finland (FI), France (FR), Germany (DE), Greece (GR), Italy



real and financial assets,<sup>3</sup> net of its total liabilities, and excludes wealth from public or occupational pension plans. The Gini coefficient of household private net wealth ranges from 0.76 for Austria to 0.56 in Greece.<sup>4</sup> The cross-country evidence on wealth inequality becomes particularly striking when it is depicted in relation to the degree of redistribution of transfers, which generally aim at reducing income inequality. Based on income data from the European Survey on Income and Living Conditions (EU-SILC) for 2004 to 2010, I first document that, within the euro zone, those countries with more generous and redistributive transfers to households robustly show higher inequality in private net wealth. The degree of redistribution of transfers is measured by the reduction in the Gini coefficient of after-tax earnings when augmenting it with public transfers.<sup>5</sup> Figure 1 shows that there is a negative correlation of -0.5.<sup>6</sup>

The redistributive effect of transfers in some countries is so substantial, that even the cross-country correlation of the Gini coefficients of after tax earnings and the Gini coefficients of net wealth switches from positive (0.12) to negative (-0.62) when adding transfers to the income measure. So interestingly, in the euro area countries where inequality in households' earnings after tax and transfers is lowest and transfers reduce the income Gini coefficient to the largest extent, private net wealth is most unequally distributed.

Since the proposed model mechanism relates wealth inequality to public insurance and, in particular, implies that there is a larger share of low-wealth households in more generous welfare states, it is important to ascertain that the differences in the wealth Gini coefficients within the euro area are not driven by differences in the share of wealth held by the richest households. Hence, the top 5th percentile of the wealth distribution in each country, owning on average about 50% of total wealth, is discarded. Importantly, their wealth accumulation process is unlikely to be affected by social insurance. The results remain robust and the correlation even increases to -0.69.<sup>7</sup> It is shown in Appendix 8 that this result is also robust to other measures of wealth inequality, such as the share of wealth

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(IT), Netherlands (NL), Portugal (PT) and Spain (ES).

<sup>3</sup>Real assets cover the household's main residence, other real estate, vehicles, valuables, and self-employment business wealth, while financial assets are composed of deposits, mutual funds, bonds, publicly traded shares, voluntary pensions etc.

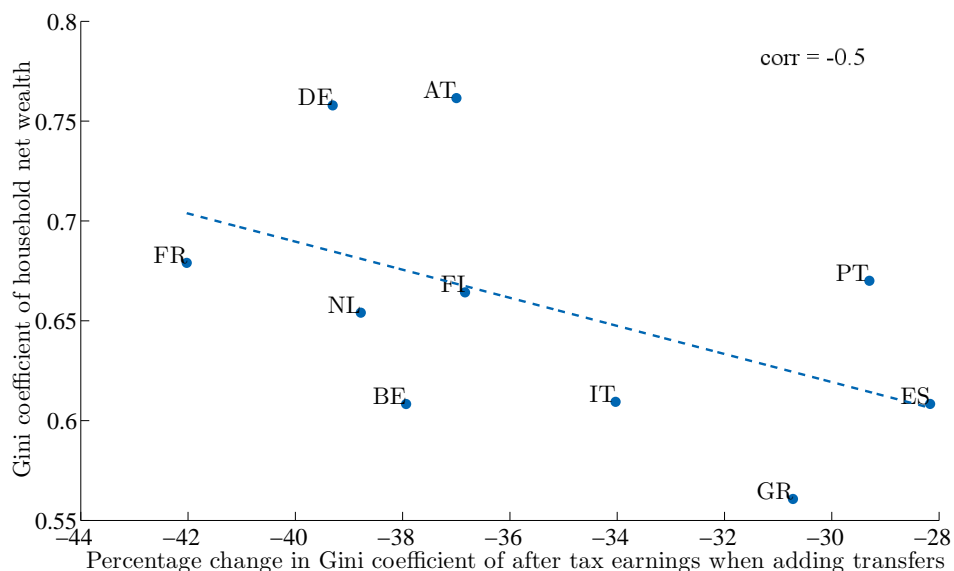
<sup>4</sup>Since the HFCS is a multiply imputed dataset, I take the average Gini coefficient over all five imputates.

<sup>5</sup>Public transfers are composed of unemployment benefits, old-age and survivors' benefits, family allowances, housing allowances, and minimum income benefits.

<sup>6</sup>In particular, this negative correlation also holds for the change in the income Gini coefficient of only working households.

<sup>7</sup>Discarding the top 10th or 20th percentile from the wealth distribution does not considerably change the relative order and level of the Gini coefficients for private net wealth and only further increases the negative correlation to -0.71 and -0.72, respectively.

Figure 1: Correlation of Gini coefficients of household net wealth and percentage change in income Gini coefficients due to transfers



Sources: *HFCS* (~2010), *Encuesta financiera de las familias españolas (EFF, 2011)*, *EU-SILC (2004-2010)*

held by the bottom 50%, excluding the top 5th percentile (see Figure 5). Accounting for cross-country differences in household compositions through equalization of net wealth also preserves the result (see Figure 6). This robustness of cross-country differences in wealth inequality with respect to the household structure has also been documented by Fessler et al. (2014) in a more profound analysis.

At first, these empirical facts are surprising, as the reduction in income differences across households through transfers would be expected to also translate into lower wealth differences. However, this paper demonstrates, in line with the argument by Hubbard et al. (1995), that generous public insurance crowds out private savings, especially of the poor, and thereby creates a larger fraction of low-wealth households. This increase in overall wealth inequality through an expansion of the left tail of the wealth distribution in the long run outweighs the mitigating effects of lower income differences through transfers in the short run, leading to higher wealth inequality in more generous welfare states.

## 3.2 Cross-country Evidence on Wealth Inequality and the Generosity of Public Transfers

In order to analyze differences in the generosity of various social security institutions in the euro zone, I consider statistics on several social security payments. Plotting those against the countries' private net wealth Gini coefficients suggests that there might be some systematic relationship between social security generosity and private net wealth inequality.

### 3.2.1 Minimum Income Support Programs

Figure 2 shows a positive correlation of 0.81 between the Gini coefficient of private net wealth (of the bottom 95%) and the absolute amount of minimum income benefits of means-tested income support programs expressed as a percentage of median net labor income of employed working-age households according to the EU-SILC. Since the EuMin database on minimum income protection in Europe (Bahle and Huble, 2012) provides the legally stipulated amounts of minimum income benefits for various household types, benefits are weighted by the household composition in the HFCS. This positive correlation is in line with the model predictions that in countries with more generous minimum income benefits, poor households have lower incentives to save for old-age and precautionary motives and hence wealth will be distributed more unequally. Notably, Italy and Greece are the only two countries which do not provide universal minimum income support to their citizens.<sup>8</sup>

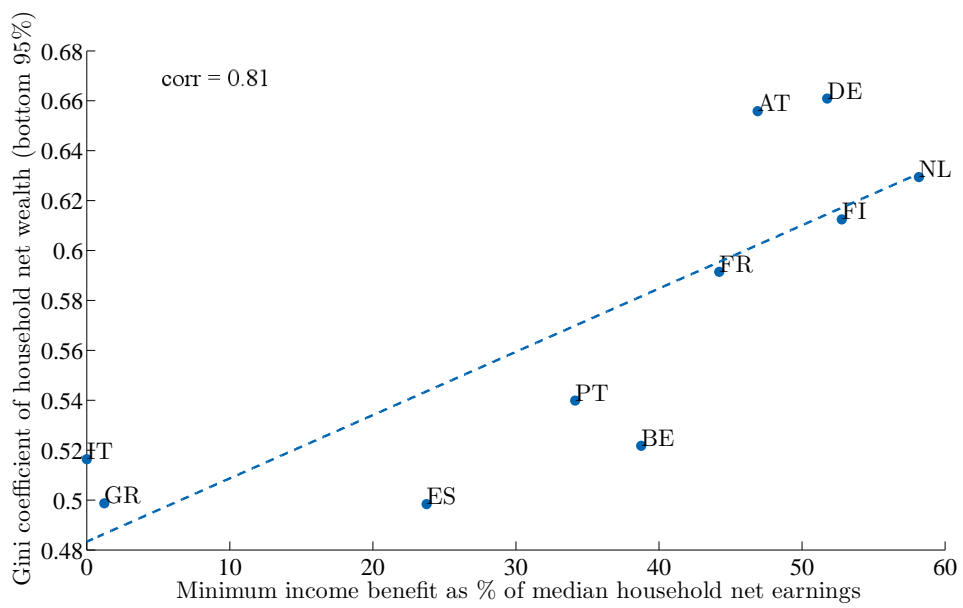
### 3.2.2 Unemployment Insurance

Furthermore, I document in Figure 3 a positive correlation of 0.45 between private net wealth inequality for the bottom 95% and the average unemployment net replacement rate over the first 60 months. Similarly to minimum income benefits, the average unemployment benefit replacement rates are weighted by the household compositions in

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<sup>8</sup>In 1998, Italy implemented local minimum income schemes in some municipalities as an experiment. However, this policy was abolished again in 2003 and replaced with an optional and poorly subsidized policy, which enabled only some wealthier regions to implement the scheme (Casas, 2005). Only in 2014, the Italian government decided to introduce a pilot project called "support for active inclusion". However, so far support is linked to previous labor market participation and focuses on families with children (Social Protection Committee, 2014). In Greece, the provision of income support is up to regional authorities and mainly targets specific groups, e.g. in old age, in mountainous regions or poor households with children. However, as can be seen in the graph, these child benefits are very low and amounted from 2004 until 2010 on average to 55 Euro per month. The main political arguments in Greece against minimum income support are budgetary constraints and high regional income diversity which hinders the government from setting one universal minimum income standard (Casas, 2005).

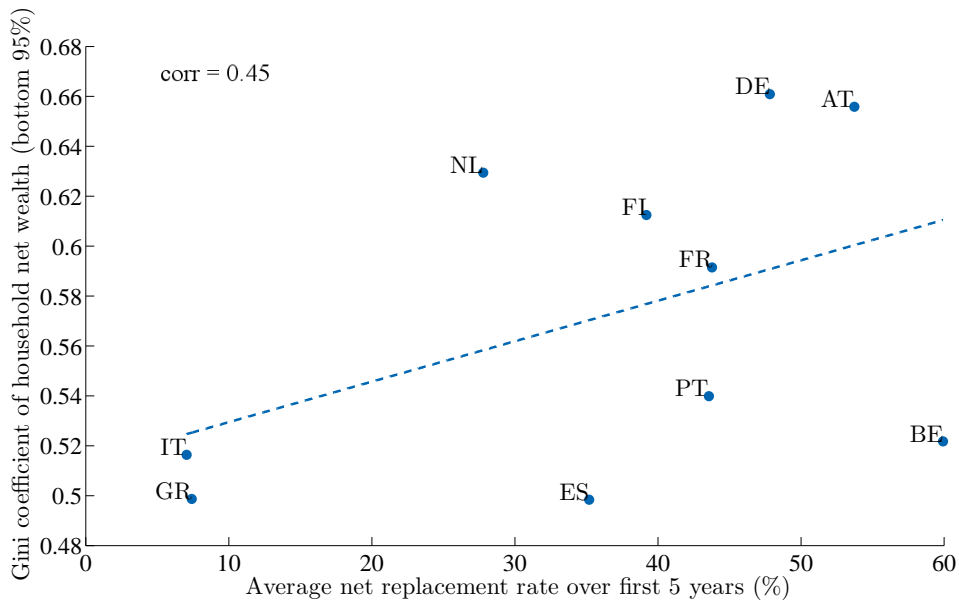
Figure 2: Wealth inequality and minimum income benefits



Sources: HFCS (~2010), EEF (2011), EuMin database/OECD "Benefits and Wages" database (2004-2008, 2010)

the HFCS. It should be kept in mind that the average net replacement rate alone is not indicative of which benefit system provides better insurance to households, since the insurance effect also depends on the duration of benefit eligibility and the expected length of the unemployment spell.

Figure 3: Wealth inequality and unemployment insurance

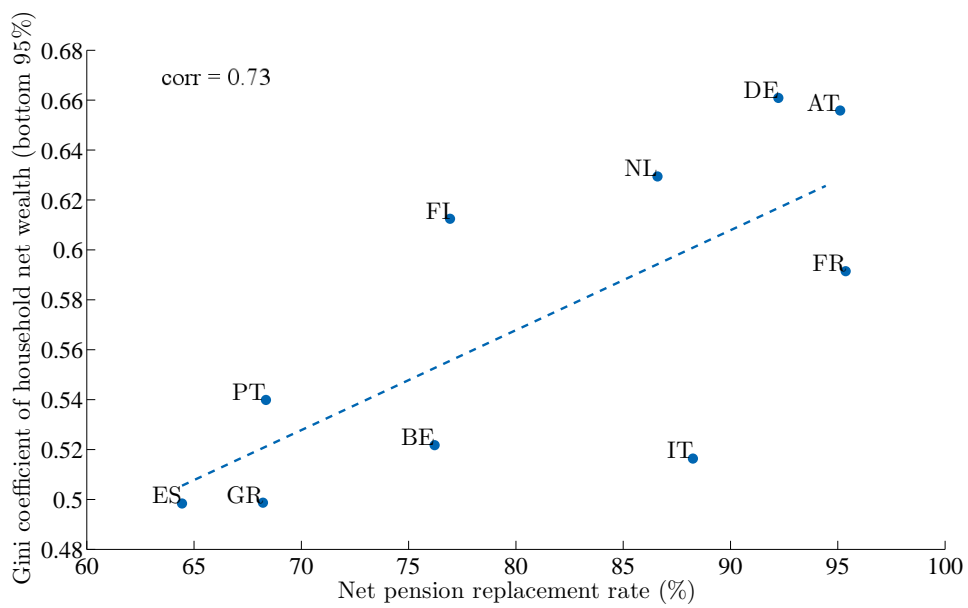


Sources: HFCS (~2010), EEF (2011), OECD “Benefits and Wages” database (2004/2010)

### 3.2.3 Public and Occupational Pension Entitlements

Figure 4 depicts cross-country differences in pension generosity and suggests a strong positive relationship of 0.73 with private net wealth inequality, a finding that is in line with the theoretical predictions of the model. The generosity of the pension system is measured by the net pension replacement rate and defined as median net old-age/survivors’ benefits of retired households aged 65 to 75 relative to median net earnings of employed and unemployed households aged 50 to 60. It is calculated from the annual waves of the EU-SILC (2004-2010). Besides public pension entitlements, it also captures survivors’ benefits and payments by occupational pension plans.

Figure 4: Wealth inequality and public pension wealth



Sources: HFCS (~2010), EEF (2011), EU-SILC (2004-2010)

## 4 Model

Following Hubbard et al. (1995), I consider a partial equilibrium life-cycle model with incomplete markets and a small open economy.

The household sector faces idiosyncratic earnings risk, i.e. stochastic fluctuations in gross earnings, but also an exogenous risk of becoming unemployed as in Wellschmied (2015). Unemployment risk is explicitly modeled to analyze the role of unemployment insurance.

Households save for old age, and for precautionary and bequest motives. The government may provide public insurance to households such as asset-based means-tested minimum income support, unemployment insurance or pension schemes. Households pay progressive labor income taxes and make social security contributions.

### 4.1 Household Income

Households' heads enter the labor market at age  $h_1$  and work for 40 years. Until retirement at age  $h_{40}$ , households' gross earnings during employment are composed of a deterministic part,  $\mu_{c,h}$ , determined by age, and a stochastic component,  $z_h$ , that captures the uncertainty and persistence of earnings shocks:

$$\omega_h = \mu_{c,h} + z_h$$

The stochastic component evolves according to an AR(1) process:

$$z_h = \rho z_{h-1} + \nu_h, \quad \nu_h \sim N(0, \sigma_c^2)$$
$$c \in \{AT, BE, FI, FR, DE, GR, IT, NL, PT, ES\}$$

Employed (e) households pay progressive labor income taxes, which also include social security contributions, and obtain net earnings:

$$w_{h,e}^{net} = (1 - \tau(\omega_h))\omega_h$$

where  $\tau_c(\omega_h)$  is a tax rate function of gross earnings. Given that a household's employment status and earnings determine its potential to accumulate wealth, a steeper age-earnings profile,  $\mu_{c,h}$ , higher gross earnings risk,  $\sigma_c^2$ , and unemployment risk also translate into higher wealth inequality. However, the effect will be muted under more progressive income taxation.

If the household becomes unemployed, the government provides unemployment insurance up to the retirement age of  $h_{40}$ . In general, unemployment benefits are modeled to replace a constant fraction of the previous period's net earnings. To avoid an additional state variable in the household's problem, previous period's net earnings are approximated by today's earnings, which would have realized if the household had not become unemployed. Since the earnings process is quite persistent, last period's net earnings are fairly well approximated.

The country-specific initial net unemployment replacement rate  $rr_c(\omega_h)$  is a function of gross earnings. Households receive benefits in the first period of unemployment with certainty and keep them in the following period with some positive probability  $p_c$ . This modeling approach is meant to capture cross-country differences in the duration of benefit eligibility.

$$b_h = \begin{cases} rr_c(\omega_h)w_{h,e}^{net}, & \text{if eligible} \\ 0, & \text{if not eligible} \end{cases}$$

Engen and Gruber (2001) have shown theoretically and empirically that higher unemployment replacement rates decrease aggregate precautionary savings. Consequently, in the case of redistributive unemployment schemes, i.e. households with lower gross income have higher replacement rates, private savings of the low-income are crowded out relatively more, thereby increasing overall wealth inequality.

However, the degree of insurance through unemployment benefits also critically depends on the overall unemployment risk, i.e. the joint-job finding and job-separation rate. Moreover, the impact of the unemployment insurance system on wealth inequality also depends on how well benefits insure households during their expected spell of unemployment. If the expected unemployment duration is relatively long or the replacement rate low, households will be incentivized to increase self-insurance and hence wealth inequality will be lower.

The household stops working at retirement age and receives public pension payments, which are a concave function  $f_c$  of the household's pre-retirement net earnings.

$$w_h^r = f_c(w_{h_{40},e}^{net}), \text{ if } h > 40$$

Consequently, the implied retirement net replacement rate declines with pre-retirement net earnings. If the household is unemployed prior to retirement, the pension payment



replaces a fraction of the net earnings that would have realized according to the stochastic process if the household had been employed in the period prior to retirement. Otherwise, pension losses due to unemployment shocks would be highly overestimated, as pensions usually replace a fraction of life-time earnings. Since earnings are calibrated to be quite persistent, pre-retirement earnings are used to proxy life-time earnings. There is no explicit retirement decision.

In my framework, there are two features of the pension system affecting wealth inequality, namely its generosity and its degree of progressivity. The more generous and redistributive the pension entitlements, the more unequally wealth should be distributed among households.

First, the model predicts a positive relationship between the generosity of pension entitlements and wealth inequality. The higher expected pension payments in the future, the less private wealth households will accumulate for old-age provision. This displacement effect does not only matter for aggregate savings, but also for wealth inequality, because households cannot access their public and occupational pension accounts during working life. The higher the public pension entitlements, the lower the overall stock of private net wealth that can be used for consumption smoothing in case of negative income shocks. Households are therefore more likely to deaccumulate assets and become borrowing-constrained in the presence of more generous pension schemes, because they can no longer pool their precautionary and old-age savings.

Second, as with unemployment benefits, a redistributive pension scheme will particularly discourage savings of low-income households (Domeij and Klein, 2002).

More generally, household net earnings before minimum income benefits are defined as

$$w_h^{net} = \begin{cases} \omega_h(1 - \tau_c(\omega_h)), & \text{if } e \text{ and } h \leq 40 \\ b_h, & \text{if } u_b \text{ and } h \leq 40 \\ \omega_{min}, & \text{if } u \text{ and } h \leq 40 \\ w_h^r, & \text{if } h > 40 \end{cases}$$

where  $e$  stands for being employed,  $u_b$  for being unemployed and eligible for unemployment benefits and  $u$  for being unemployed and no longer eligible.  $\omega_{min}$  is meant to capture a minimum income that can be privately obtained by the household when un-

employed, e.g. through private transfers from family members outside the household.

If unemployment benefits expire or the working income is too low to cover basic household expenses, households may become eligible for minimum income benefits, i.e. the government guarantees a minimum consumption floor,  $T\bar{R}_c$ . Minimum income support programs are considered to be households' last public safety net. In contrast to unemployment benefits and pension payments, they are universal and hence do not depend on past contributions, but only on the households' current means in terms of assets and income. Therefore, the actual transfer,  $TR_c$ , made to the household negatively depends on its choice of end-of-period wealth,  $k_h$ , and its net income,  $w_h^{net}$ , and is zero, if both exceed  $T\bar{R}_c$ .

$$TR_c(k_h, w_h^{net}) = \max\{0, T\bar{R}_c - k_h(1+r) - w_h^{net}\}$$

Asset-tested minimum income benefits were first introduced into a life-cycle model in a seminal paper by Hubbard et al. (1995). The effect of means-tested minimum income benefits on wealth inequality is twofold. First, minimum income support reduces, in particular, the downward income risk for low-income households and thereby lowers their need to self-insure to a relatively greater extent. They show that households with high expected life-time income still maintain the usual incentives to save for precautionary purposes, while this motive is highly distorted for the lower part of the income distribution for whom minimum income benefits make up a larger fraction of their life-time income, thus amplifying the wealth gap between high and low-income households. Second, asset-based means-testing introduces an implicit tax on savings and hence households face a trade-off between saving for precautionary motives and dissaving to become eligible for income support.

## 4.2 Household Optimization Problem

Each period, the household chooses its total consumption,  $c$ , and end-of-period assets,  $k$ , given its beginning-of-period assets,  $a$ , labor income,  $z$ , and employment status. In particular, it forms expectations about whether it will be alive and employed in the next period and if so, what the resulting labor market income will be. In case the household was employed the previous period and becomes unemployed, it will always receive unemployment benefits in the first year. The dynamic planning problem of the household subject to the budget and non-negativity constraint will be presented for each labor market status  $\{e, u_b, u\}$ .

The Bellman equation of the employed household at age  $h$  is:

$$V(h, a, z, e) = \max_{c, k} \left\{ u(c) + \beta \mathbf{E} \left\{ (1 - \iota_h) [(1 - \delta_c) V(h + 1, a', z', e) + \delta_c V(h + 1, a', z', u_b)] + \iota_h \phi(a') \right\} \right\}$$

$$\begin{aligned} s.t. : c + \frac{a'}{1+r} &= a + w^{net} + \frac{TR_c(k, w^{net})}{1+r} \\ a' &= (1+r)k + TR_c(k, w^{net}) \geq 0 \end{aligned}$$

where  $\delta_c$  is the country-specific probability of job separation and  $\mathbf{E}$  is the expectation operator. The household dies with probability  $\iota_h$  at age  $h$  and with certainty at the age of 84. For every deceased household, a new household will be born. Beginning-of-period assets in the next period,  $a'$ , correspond to the end-of-period asset choice of the household, the earned dividend and potential end-of-period transfers, and have to be non-negative. The fact that low net wealth households close to the borrowing constraint are likely to be income poor households with limited access to credit, motivates the assumption of zero borrowing across all countries. The utility function displays constant relative risk aversion (CRRA) with risk aversion parameter  $\xi$ :

$$u(u) = \frac{1}{1-\xi} u^{1-\xi}, \quad \xi > 0,$$

For the bequest function, I choose the specification as in De Nardi and Yang (2015):

$$\phi(a) = \phi_1 \frac{(a + \phi_2)^{1-\xi}}{1-\xi}, \quad \xi > 0,$$

The parameter  $\phi_1$  governs the desire to leave bequests, while the parameter  $\phi_2$  reflects the extent to which bequests are luxury goods. Note that households' preferences for leaving bequests are assumed to be equal across all countries, as the analysis of cross-country differences in the bequest motives are beyond the scope of this paper.

The optimization problem of a currently unemployed household receiving benefits is:

$$\begin{aligned}
V(h, a, z, u_b) = \max_{c,k} & \left\{ u(c) + \beta \mathbf{E} \left\{ (1 - \iota_h) \left[ \gamma_c V(h+1, a', z', e) \right. \right. \right. \\
& + (1 - \gamma_c) [p_c V(h+1, a', z', u_b) + (1 - p_c) V(h+1, a', z', u)] \\
& \left. \left. \left. + \iota_h \phi(a') \right\} \right\} \\
s.t. : c + \frac{a'}{1+r} & = a + b + \frac{TR_c(k, b)}{1+r} \\
a' & = (1+r)k + TR_c(k, w^{net}) \geq 0
\end{aligned}$$

The household either finds a new job with probability  $\gamma_c$  or stays in unemployment, in which case it keeps its benefits with probability  $p_c$ . This modeling approach is meant to capture cross-country differences in the duration of benefit eligibility. I also allow for cross-country differences in unemployment rates through country-specific parameters for the job separation rate,  $\delta_c$ , and job finding rate,  $\gamma_c$ , as the ultimate effect of the unemployment insurance system on wealth inequality also depends on how well benefits insure households during their expected period of unemployment.

The optimization problem of a currently unemployed household no longer receiving unemployment benefits is:

$$\begin{aligned}
V(h, a, z, u) = \max_{c,k} & \left\{ u(c) + \beta \mathbf{E} \left\{ (1 - \iota_h) \left[ \gamma_c EV(h+1, a', z', e) \right. \right. \right. \\
& \left. \left. \left. + (1 - \gamma_c) V(h+1, a', z', u) \right] + \iota_h \phi(a') \right\} \right\} \\
s.t. : c + \frac{a'}{1+r} & = a + \omega_{min} + \frac{TR_c(k, \omega_{min})}{1+r} \\
a' & = (1+r)k + TR_c(k, \omega_{min}) \geq 0
\end{aligned}$$

In the present paper, I allow for a simple “warm-glow” type of bequest motive to ensure that not all households deaccumulate their wealth during retirement. In order to analyze the role of bequests for euro area differences in wealth inequality, and in particular the interaction between bequests and social security provision, a more enhanced model of bequests would be needed, where young households anticipate and inherit only the wealth accumulated and left by their parents. In this kind of framework, bequests

generally lead to an amplification of wealth inequality in the presence of public insurance, since social security “disinherits” the poor (Gokhale et al., 2001). However, allowing for this type of bequest here would substantially increase the computational burden of the model, because young households also would have to take into account their parents’ state variables in order to form expectations about the size of their future bequest. Therefore, the analysis of the role of bequests for cross-country wealth inequality differences is left for future work.

## 5 Calibration

### 5.1 Household Parameters

The model’s parameters are calibrated at an annual frequency and the baseline parameters are reported in Table 1. Despite the fact that each parameter is presented to be calibrated individually to match a specific data moment, it has to be kept in mind that all parameters are of course calibrated jointly.

For the felicity and bequest function, I calibrate the coefficient of relative risk aversion to a value of  $\xi = 1.5$ , as in a closely related paper on minimum income benefits by Wellschmied (2015). The two parameters  $\phi_1$  and  $\phi_2$  governing the bequest motive are pinned down by matching the average and median wealth of households at the end of the life cycle relative to average and median wealth of younger households in the euro area.  $\phi_1$  determines the overall strength of the bequest motive, and is chosen to match the average across all euro area countries’ ratios of mean wealth of households older or equal to 84 years relative to households younger than 84 of 0.56 (see Table 8 Appendix 10).  $\phi_2$  mainly affects the distribution of wealth at older ages, as for  $\phi_2 > 0$  bequests become a luxury good and only households at a certain threshold of wealth will have the desire to leave bequests. Hence, I match in the model, as a second moment, the average across all countries’ ratios of median wealth of households older or equal to 84 years relative to median wealth of their younger counterpart.

The annual real interest rate is set to 2.5%, which corresponds to the average across the countries’ annual yields of 10-year national government bonds traded in the secondary market and is adjusted for inflation as measured by the annual rate of change of the Harmonised Index of Consumer Prices.<sup>9</sup> The choice of this specific annual real interest rate is however not crucial for the model’s implied cross-country differences in wealth

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<sup>9</sup>Government bond yields and inflation rates cover the years from 1997 to 2010 and are provided by the ECB Statistical Data Warehouse.

inequality, as e.g. a higher real interest rate of 4% would leave the explanatory power of the model unchanged. The model's time preference parameter,  $\beta$ , is set to the same value of 0.983 across all ten euro area countries ( $N=10$ ) and chosen to equalize the average over all countries' Gini coefficients in the data and in the model:

$$\bar{\lambda} = \frac{\sum_{c=1}^N \lambda_c}{N} = \frac{\sum_{c=1}^N \hat{\lambda}_c(\beta, \theta_c)}{N}$$

where  $\lambda_c$  is the Gini coefficient of private net wealth in country  $c$  according to the HFCS, after discarding the top 5th percentile from the wealth distribution in each country. The model's predicted wealth Gini coefficient,  $\hat{\lambda}_c(\theta_c)$ , for country  $c$  is a function of  $M$  country-specific parameters,  $\theta_c = \{\theta_c^1, \dots, \theta_c^M\}$ , describing its labor market process and welfare policies. In the baseline analysis, I choose to discard the top 5th percentiles from the actual wealth distributions for the calculation of  $\lambda_c$ , as the model is, like most incomplete markets models, incapable of matching the high wealth inequality levels observed in the data without generating an overly large fraction of zero wealth households.

Table 1: Calibrated Parameters

Parameter		Description	Target/Source
$\beta$	0.983	Discount factor	Mean over countries' Gini coefficients of private net wealth of 0.573 (HFCS, ~2010)
$\xi$	1.5	Coefficient of RRA	Wellschmied (2015)
$r$	2.5%	Annual real interest rate	Mean over countries' real annual yields of 10-year government bonds traded on secondary market (ECB SDW, 1999-2010)
$\phi_1$	28	Bequest utility	Avg. euro area ratio of median/mean wealth of heads aged $\geq 84$ relative median/mean wealth of heads aged $< 84$ : 0.3167/0.5559
$\phi_2$	8.8	Bequest utility shifter	
$a_{h_1}$	€5813	Initial asset level at age 22	Mean over countries' median asset holdings of households with heads aged 20-25
$\omega_{min}$	0	Private transfers to unemployed household	Median net private transfers received by unemployed households (EU-SILC, 2004-2010)
$\iota_h$		Probability of dying	Life tables for euro area countries, Eurostat (1995-2010)
$h_1$	22	Age of labor market entry	40 years of working life
$h_{40}$	62	Age of retirement	Avg. pensionable age in (OECD, 1999-2010)
$h_{62}$	84	Age of decease	Oldest age in life table

All countries are assumed to have the same annual survival probabilities for house-

holds' heads at a given age. They are set to the average across all country-specific survival probability rates. This data is provided by the Eurostat life tables for the time span from 1995 to 2010. The survival probability of households at age 84 is set to zero. To account for the current age structure of the population in euro area, I average across the countries' age distributions in the HFCS, approximate the resulting euro area age distribution with a polynomial function of order five and apply the obtained age weights to the model before computing the Gini coefficients.

Households at the age of 22 start with an initial asset level of  $a_{h_1} = \text{€}5813$ , which corresponds to the mean over countries' median asset holdings of households with heads aged between 20 and 25.<sup>10</sup> The median unemployed household in the EU-SILC received zero net private transfers from 2004 to 2010. To test for robustness, I will also allow for positive private transfers in section 6.

I then assume that all countries share the same baseline parameters shown in Table 1 and I allow them to differ only in the parameters describing the unemployment and net earnings process, as well as the social security system. This way, I can examine the variation in cross-country wealth inequalities that are generated by the countries' specificities of their social security institutions and labor markets.

## 5.2 Labor Market Process Before Retirement

Earnings and unemployment dynamics are assumed to be purely exogenous. The job separation and finding rates are chosen to match the average unemployment rate and percentage share of long-term unemployed in each country. The household unemployment rates for the time span 2004 to 2010 in the ten euro area countries are calculated from the EU-SILC dataset using weights and are displayed in Table 2. They are derived from the employment status of the household's head who is identified as the household member with the highest personal income, in terms of gross earnings and public individual transfers such as e.g. unemployment benefits. The percentage of long-term unemployed households, i.e. households with heads being unemployed for more than a year, cannot be determined from the EU-SILC directly and is hence approximated with the average percentage share of long-term unemployed individuals reported by Eurostat from 2004 to 2010.

The earnings risk is estimated for each country from EU-SILC gross earnings data of employed households' heads between 25 and 60 years. I assume that the observed log-income of a household,  $y_{i,h,t}$  is composed of a deterministic part,  $f(o_{i,h,t})$ , determined

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<sup>10</sup>Note that households' initial wealth does not necessarily coincide with the amount of bequests left by deceased households.

by observable household characteristics of the household head,  $o_{i,h,t}$ , and a stochastic component,  $y_{i,h,t}^*$ , which follows an AR(1) process with persistence  $\rho$ :

$$y_{i,h,t} = f(o_{i,h,t}) + y_{i,h,t}^*, \quad (1)$$

$$y_{i,h,t}^* = \rho y_{i,h-1,t-1}^* + \nu_{i,h,t}, \quad (2)$$

$$\nu_{i,h,t} \sim N(0, \hat{\sigma}_c^2) \quad (3)$$

where  $t$  is the year and  $h$  is the age of household  $i$ . Earnings shocks,  $\nu_{i,h,t}$ , are assumed to be drawn from a log-normal distribution. This log-normality assumption, although it usually does not hold perfectly, allows me to approximate the earnings process using a Markov chain with seven income states.

I estimate the deterministic component,  $f(o_{i,h,t})$  by regressing log-earnings on an age polynomial of order 4, education and gender dummies, and the household composition. In order to control for variations in household size and composition over the life cycle, I include the number of heads (single or couple), the number of children younger than 18 and the number of other dependent adults in each household. I eliminate any observation where the residual of this regression,  $y_{i,h,t}^*$ , belongs to the bottom or top 0.5 % of all residuals.

Following Hintermaier and Koeniger (2016), the sample variance of the residuals,  $y_{i,h,t}^*$ , from this regression is used to derive the country-specific short-run variance of gross earnings shocks,  $\hat{\sigma}_c^2$ , assuming the same persistence of  $\rho = 0.95$  of earnings shocks across all countries, a common estimate in the empirical literature. This parameter is calibrated rather than estimated, such that the cross-country differences predicted by the model do not depend on imprecise estimates of this parameter, given the short time horizon of only 3 years available for estimation in some countries.<sup>11</sup> While controlling for household characteristics leads to an underestimation of overall earnings heterogeneity in the model, it allows me to obtain a better proxy for pure earnings risk. This is important because the degree of earnings risk matters for the accumulation of precautionary savings and its distribution, which aside from savings for old age is of main interest in this paper.

Since the EU-SILC panel only follows a household for three consecutive years, it is not possible to observe the full life-cycle earnings of a cohort. However, I make use of the cross-sectional age-earnings patterns to approximate deterministic life-cycle profiles,  $\mu_{c,h}$ , in the respective countries. Therefore, I determine median gross earnings at each age and smooth the profile using a fourth order polynomial. Since, due to positive real earnings

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<sup>11</sup>The EU-SILC Survey was implemented in 2004 in Austria, Belgium, Finland and France, in 2005 in Germany and the Netherlands, in 2006 in Spain, and in 2007 in Greece, Italy and Portugal.



growth e.g. cohorts at the age of 30 have on average a higher nominal median earnings level than cohorts of currently 50-years did have 20 years ago, this cross-sectional age-earnings profile underestimates earnings growth over the life cycle. Hence, in order to make average earnings comparable across cohorts, I need to transform median earnings at a given age relative to the base age of 50 by multiplying it with the real wage growth factor  $(1 + g)^{age-50}$ . I adjust all earnings profiles with the euro area average real growth rate of 1% calculated from average real wages from 1990 until 2010 provided by the OECD Database.<sup>12</sup> Figure 7 shows the countries' cross-sectional age-earnings profiles, adjusted for real euro area wage growth. Furthermore, they are divided by the average Purchasing Power Parity index from 2004 to 2010 of the respective country to ensure comparability across the euro area. Note that the results are also robust to using average gross earnings instead of median earnings for the calibration of the age-earnings profile.

As the EU-SILC is a survey specifically designed to measure household income, gross earnings data from the EU-SILC is preferred over data from the HFCS. While households interviewed in the EU-SILC are explicitly asked for all potential income sources, earnings questions in the HFCS are more broadly categorized, thereby increasing the risk of imprecise measurement.

## 5.3 Policy Parameters

### 5.3.1 Minimum Income Support

Information on the absolute amount of minimum income benefits is taken from a comparative database on minimum income protection in Europe, which provides annual data from 2004 until 2008. The data is based on the OECD tax and benefit model which simulates minimum income benefits for various household types (single person or married couple, without children or with 2 children). However, the corresponding OECD "Benefits and Wages" database only provides information for the years 2005, 2007 and 2010. Therefore, I make use of the data provided by the EuMin database until 2008 and complement it with OECD data for 2010 and interpolate in-between for 2009. Minimum income benefits cover cash benefits, including housing benefits as well as child benefits. Since e.g. married couples with children are entitled to more generous social assistance, I take the different household compositions in the euro area countries into account when

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<sup>12</sup>Average real wages are obtained by dividing the national-accounts-based total wage bill by the average number of employees in the total economy, which is then multiplied by the ratio of the average usual weekly hours per full-time employee to the average usual weekly hours for all employees. They are measured in USD constant prices using 2012 as a base year and Purchasing Power Parities for private consumption in the same year.

Table 2: Parameters of labor market risk

Country	Std. of earnings shocks $\sigma$	Job sep. rate $\delta$	Job find. rate $\gamma$	Unempl. rate	Fract. of long-term unemployed ( $>1$ y)
AT	0.1809	0.0351	0.7450	6.4%	53%
BE	0.1501	0.0405	0.4590	9.8%	23%
FI	0.1539	0.0840	0.7560	8.3%	30%
FR	0.1599	0.0611	0.6100	7.8%	23%
DE	0.1802	0.0447	0.4810	10.5%	40%
GR	0.1885	0.0578	0.5090	5.6%	48%
IT	0.1723	0.0417	0.4220	6.7%	49%
NL	0.1533	0.0370	0.6610	2.9%	48%
PT	0.1758	0.0437	0.5550	7.6%	49%
ES	0.1750	0.1021	0.5880	9.9%	27%

computing the average expected entitlements by weighting with the corresponding percentage of households of each type in the sample.<sup>13</sup> Figure 2 shows the weighted absolute amount of minimum income benefits expressed as a percentage of median net earnings of employed households aged 25 to 60 in the EU-SILC sample.<sup>14</sup>

<sup>13</sup>Using information on the relationship and age of household members in the HFCS, I assign all households in each country to 4 different types which are meant to approximate the aforementioned types stipulated in the OECD benefit model: Single head or head in partnership/marriage, without children or with at least one child.

<sup>14</sup>Spain constitutes a special case, as it is the only country where minimum income provision is a regional competence and conditions of payment can hence vary across regions. Therefore, the minimum income support for Spain reported in the EuMin database only refers to the amount of minimum income benefits available to households resident in the community of Madrid. Moreover, minimum income benefits are only of unlimited duration in the six autonomous communities of Asturias, Castilla y Leon, Madrid, Cataluna, Extremadura and Valencia (length of 3 years), which were inhabited by 50.7% of Spain's total population in 2011. The calibration strategy for Spain is to solve the model twice, first for regions for which the model assumes no minimum income scheme for simplicity, and second, for the remaining regions under the assumption that these provide unlimited minimum income benefits of an amount equal to that of Madrid. The overall implied Gini coefficient of private net wealth in Spain is computed from the weighted average of the two resulting wealth distributions.

### 5.3.2 Unemployment Insurance

The OECD database “Benefits and Wages” reports the initial unemployment net replacement rates for multiples of average worker (AW) gross earnings and 6 different household types (single person, one-earner married couple or two-earner married couple, without children or with two children). Since unemployment net replacement rates differ for distinct household types due to e.g. potential family, childcare or lone-parent benefits, the average net replacement rate used for the calibration of the model is obtained by weighting with the corresponding fraction of households of each approximate type in the sample. Similarly to the weighting of minimum income benefits, I use information from the HFCS on age, and employment status and relationships of household members, to assign households to 6 categories. Since unemployment benefits depend on an additional characteristic of the household, namely whether the household is a one-earner or two-earner married couple, I also use information on the employment status to classify households.

Figure 8 in Appendix 10 depicts the initial net replacement rate, averaged over the years 2004 and 2010, as a function of multiples of average gross earnings and reveals that unemployment benefits replace a larger fraction of previous earnings for low-income households.<sup>15</sup> The country-specific probability,  $p_c$ , of keeping benefits from the second year on is chosen to match the average net replacement rates over the first 5 years of unemployment, a statistic also provided by the OECD database.

### 5.3.3 Public and Occupational Pension Scheme

The net pension replacement rate as shown in Figure 4 is defined as the median net old-age and survivors’ benefits of retired households aged 65 to 75 relative to median net earnings of employed and unemployed households aged 50 to 60 in the EU-SILC. Since most occupational pension plans held by households are still of type defined benefit and future payments are therefore dependent on unknown future conditions, the HFCS measures households’ entitlements to their occupational pension plans relatively poorly. Therefore, I follow the Household Finance and Consumption Network and only include private pension wealth in the calculation of total private net wealth. Instead, the net pension replacement rate includes income from both public and occupational pension plans, but excludes pension income from individual private plans. In order to allow

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<sup>15</sup>For the model’s calibration, I assume that households with previous gross earnings levels smaller than 67% of average gross earnings have the same net replacement rate as households with previous earnings equal to 67% of average gross earnings. The net replacement rates of households with previous gross earnings larger than 1.5 times of average gross earnings are extrapolated and stay constant.

Table 3: Policy Parameters

Parameter	Target	Data Source
<b>Minimum income benefits</b>		
$T\bar{R}_c$	Income floor (weighted by household composition)	EuMin database (2004-2008), OECD “Benefits and Wages” database (2010)
<b>Unemployment insurance</b>		
$rr_c(\omega_h)$	Initial net replacement rate (weighted)	OECD “Benefits and Wages” database (2004/2010)
$p_c$	Avg. net replacement rate over first 5 years of unemployment	
<b>Pensions</b>		
$f_c(x) = a_c x^{\frac{1}{b_c}}$	Median net pension replacement rate	EU-SILC (2004-2010), adjusted for real wage growth
	Pension progressivity index $1 - \frac{gini(w_{41}^r)}{gini(w_{40}^{net})}$	
<b>Earnings tax</b>		
$\tau_c(\omega_h)$	Avg. tax rate schedule defined on multiples of average earnings	OECD tax database (2010)

for comparability of incomes of different cohorts, I adjust earnings and pensions for the average euro area real wage growth of 1%, as described in Section 5.2. In order to capture the extent to which the public and occupational pension systems are redistributive, the pension progressivity index is calculated. In the model, the index is defined as 1 minus the ratio of the Gini coefficient of net pension income  $w_{41}^r$  relative to the Gini coefficient of pre-retirement net income of employed and unemployed households  $w_{40}^{net}$  (see formula in Table 3). For this measure, I again refer to the two age groups used for the calculation of the net replacement rate. If pensions are perfectly proportional to pre-retirement income and hence not redistributive, the Gini coefficient of pensions is equal to the Gini coefficient of pre-retirement earnings and the progressivity index corresponds to 0. If there were a flat-rate pension scheme instead, the Gini coefficient of pensions is zero and the index would take a value of 1. I assume that pension payments are a concave function of the household’s pre-retirement net earnings:  $f_c(x) = a_c x^{\frac{1}{b_c}}$ . I choose the constant  $a_c$  and  $b_c$  in order to match the pension progressivity index and median net replacement rate in the EU-SILC. The resulting net replacement rates are declining convex functions of pre-retirement net earnings and are shown in Figure 9 of Appendix 10. For the calibration, I choose an empirical estimate over the theoretical net replacement rates provided by the OECD, because the latter neglect public pensions paid to state employees, which are

quite generous in Germany and France.<sup>16</sup>

### 5.3.4 Labor Income Taxes and Social Security Contributions

The OECD income tax database provides the average labor income tax rates for various levels of gross earnings. It includes central and sub-central government income taxes as well as employee social security contributions. The tax schedule which is defined over multiples of average earnings is directly applied in the model and interpolated in-between. Negative average tax rates are ruled out. Since the OECD tax database only provides the average tax rates for up to twice the average earnings, top marginal tax rates are used to infer the average tax rate for higher earnings (see Guvenen et al. (2014) for the applied method).

## 6 Results

This section presents the main results. In the following, I seek to answer which fraction of euro area variation in wealth inequality can be attributed to differences in labor market dynamics and the various features of the social security system. Furthermore, I will identify how much each factor contributes individually to this fraction.

### 6.1 Quantitative Importance of Labor Income Dynamics and Welfare Policies

First, I will determine which fraction of euro area variation in wealth inequality can be explained by all factors of interest. In order to determine the predictive power of the model for cross-country differences in wealth inequality, I make use of the coefficient of determination. It is a common measure to determine the goodness of fit of forecasting models. The model's predicted wealth Gini coefficient,  $\hat{\lambda}_c(\boldsymbol{\theta}_c)$ , for country  $c$  is based on a parameter vector of all the country-specific parameters,  $\boldsymbol{\theta}_c$ , describing its welfare policies and labor market process, and can be interpreted as a forecast of the actual wealth Gini coefficient,  $\lambda_c$ . Let  $\hat{\epsilon}_c$  denote the forecast error of the model when predicting the wealth Gini coefficient of country  $c$ :

$$\hat{\epsilon}_c = \lambda_c - \hat{\lambda}_c(\boldsymbol{\theta}_c)$$

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<sup>16</sup>The OECD replacement rates measure the theoretical net pension replacement rate of a representative worker who works a full career and enters the labor market today taking into account all to date enacted pensions reforms. So while the OECD net replacement rates refers to individuals, the EU-SILC median net replacement rate refers to a household unit. Furthermore, the OECD theoretical replacement rate might overestimate pension entitlements in countries with higher unemployment rates or lower labor market participation rates.

The model's time preference parameter,  $\beta$ , is set equally across all ten countries ( $N=10$ ) and calibrated to equalize the average across all country's Gini coefficients in the data and in the model, given the country-specific parameterization. This implies an average model prediction error,  $\bar{\hat{\epsilon}}$ , of zero and hence no systematic over- or underestimation of wealth inequality levels for the euro area countries considered. Assuming that all the countries share the same baseline parameter values shown in Table 1, I allow them to differ only in the parameters describing the unemployment and net earnings process, as well as the social security system. The coefficient of determination is used to quantify the predictive power of the model for cross-country variations in wealth inequality. The coefficient of determination,  $R^2$ , relates the total sum of squared forecast errors generated by the calibrated model to the total sum of squared forecast errors implied by a benchmark model predicting the same Gini coefficient, namely the mean, for each country. However, squaring forecast errors leads to an unequal weighting of small and large forecast errors. Therefore, in order to equally weight each country's forecast for the overall assessment of the model, a modified coefficient of determination,  $R$ , is introduced which expresses forecast errors in absolute instead of squared terms. It is defined as:

$$R = 1 - \frac{\sum_{c=1}^N |\hat{\epsilon}_c|}{\sum_{c=1}^N |\lambda_c - \bar{\lambda}|}$$

Relative to other constants, the mean is the most suitable benchmark forecast to evaluate the model's predictive power due to its property of minimizing the sum of absolute forecast errors:

$$\bar{\lambda} \in \operatorname{argmin} \left\{ \sum_{c=1}^N |\lambda_c - x| \right\}$$

Table 4 quantifies the overall importance of welfare policies and labor income dynamics for cross-country variations in wealth inequality using  $R$ . Overall, the model results suggest that those factors can explain 70.1% of the differences in wealth inequality across the euro area for the bottom 95% in each country's wealth distribution. Importantly, note that the parameter vector is not chosen to maximize this statistic, but is calibrated to the observed differences in welfare policies and labor market dynamics across countries. As it turns out, the modified measure,  $R$ , is more conservative compared to  $R^2$ , which implies for the same predictions an explanatory power of 89.2%. The higher  $R^2$  originates from the fact that the model performs particularly well in forecasting the large differences in wealth Gini coefficients across countries.

The first two columns of Table 4 report the 95%-wealth Gini coefficient for all countries according to the HFCS, first in levels, in column (a), and then in column (b) and (c)

expressed as a deviation and squared deviation from the mean across all euro area Gini coefficients. Column (d) and (e) depict the same statistics as column (a) and (b) for the model predictions,  $\hat{\lambda}_c(\beta, \theta_c)$ . Negative deviations imply that the wealth Gini coefficient in the respective country is below average, while positive deviations indicate countries in the euro area with above-average wealth inequality. Comparing the signs of the deviations in column (b) and (e) reveals that for every country the model correctly predicts the relative ranking of wealth inequality with respect to the mean. Furthermore, the table shows in column (f) the prediction errors of the model's forecasts, as well as absolute errors in column (g). The row labeled "Mean" in Table 4 demonstrates in column (a) and (d) that the average of Gini coefficients  $\lambda_c$  in the data of 0.573 equals, through the calibration of  $\beta$ , the mean of the model's implied Gini coefficients. For the same reason, the average forecast error of the model in column (f) is zero.

In the last row of the table, labeled  $R$ , the modified coefficient of determination is reported. It indicates that the model can explain 70.1% ( $= 1 - \frac{0.0172}{0.0575}$ ) of the cross-country differences in wealth inequality for the bottom 95% of the private net wealth distributions in 2010. Column (h) reports how well the model performs for the respective countries and the explanatory power of the model ranges from 37.4% for France to 97.5% for Italy. When including the richest 5% in the calculation of the Gini coefficients, the explanatory power of the model drops to 27.8% after recalibrating the preference parameter to  $\beta = 0.961$  in order to match the average across all country's wealth Gini coefficients of 0.654 in the data. This finding is consistent with the notion that the wealth accumulation process of the wealthiest 5% is unlikely to be driven by one of the savings motives considered in this model.

Table 4: Gini coefficients of private net wealth in the data and model

	bottom 95%										all	
	Data					Model					Forecast Error	Goodness of fit
	$\lambda_c$	$\lambda_c - \bar{\lambda}$	$ \lambda_c - \bar{\lambda} $	$\hat{\lambda}_c(\beta, \theta_c)$	$\hat{\lambda}_c(\beta, \theta_c) - \bar{\lambda}$	$\hat{\lambda}_c$	$ \hat{\epsilon}_c $	$1 - \frac{(g)}{(c)}$	$\hat{\epsilon}_c$	$ \hat{\epsilon}_c $	$1 - \frac{(g)}{(c)}$	
	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)	(i)	(j)	(k)	
Austria	0.656	-0.083	0.0833	0.634	-0.061	-0.022	0.0219	73.8%	-0.022	0.0219	73.8%	
Belgium	0.522	0.051	0.0508	0.535	0.037	0.013	0.0135	73.5%	0.013	0.0135	73.5%	
Finland	0.612	-0.040	0.0399	0.608	-0.035	-0.005	0.0048	88.1%	-0.005	0.0048	88.1%	
France	0.591	-0.019	0.0190	0.603	-0.031	0.012	0.0119	37.4%	0.012	0.0119	37.4%	
Germany	0.661	-0.088	0.0884	0.624	-0.052	-0.037	0.0366	58.6%	-0.037	0.0366	58.6%	
Greece	0.499	0.074	0.0738	0.525	0.048	0.026	0.0262	64.5%	0.026	0.0262	64.5%	
Italy	0.516	0.056	0.0561	0.518	0.055	0.001	0.0014	97.5%	0.001	0.0014	97.5%	
Netherlands	0.629	-0.057	0.0569	0.607	-0.034	-0.023	0.0227	60.2%	-0.023	0.0227	60.2%	
Portugal	0.540	0.033	0.0326	0.547	0.025	0.007	0.0073	77.5%	0.007	0.0073	77.5%	
Spain	0.498	0.074	0.0741	0.524	0.048	0.026	0.0256	65.4%	0.026	0.0256	65.4%	
<b>Mean</b>	0.573	0.000	0.0575	0.573	0.000	0.000	0.0172		0.000	0.0172		
<b>R</b>								<b>70.1%</b>			<b>27.8%</b>	



## 6.2 Decomposition

This section introduces a decomposition method which quantifies the separate effects of labor market dynamics, the pension system, unemployment insurance and minimum income benefits for the fraction explained by the model of 70.1%, as reported in Table 4. The decomposition method presented here is most closely related to one adopted in a paper by Guvenen et al. (2014) in which they analyze the extent to which differences in wage inequality between the US, their chosen reference country, and six other central European countries can be attributed to differences in labor income tax progressivity and also provide a decomposition.

For the decomposition exercise, I construct a fictive euro area country (EA) that serves as a benchmark country. Its policy parameters are defined as the average across all country-specific individual parameters,  $\theta_c^i$ :

$$\theta_{EA}^i = \frac{\sum_{c=1}^N \theta_c^i}{N}, \quad i = 1, \dots, M$$

By sequentially setting the country-specific parameters to those of the reference country and recalibrating the discount factor,  $\beta$ , to match the mean wealth Gini coefficient  $\bar{\lambda}$  in the data, the contribution of the income process and each policy to the euro area variation in wealth inequality can be determined. Finally, when setting all parameters to those of the fictive euro area benchmark country, all countries will exhibit the same mean Gini coefficient,  $\lambda_{EA}(\theta_{EA}) = \bar{\lambda}$ , as no differences remain.

Table 5 demonstrates the decomposition method which disentangles the contribution of each factor to the explanatory power of the model for cross-country wealth inequality differences. First, in column (1), I set the income process and all welfare policies to the country-specific parameters and obtain  $R$  as reported in Table 4.

In the next column (2), I then assume that all countries have the same labor income and unemployment dynamics as in the EA reference country, but still differ in the social security dimensions considered. At this stage the discount factor,  $\beta$ , is recalibrated in order to equalize the average of actual and predicted wealth Gini coefficients across the euro area and  $R$  is again determined. This step allows me to separate the role of those three policies from that of the labor market dynamics. The difference of the coefficient of determination,  $R$ , in columns (1) and (2) in the last row provides a useful measure of the role of the income process, which accounts for 12.2% (= 70.1% – 57.9%) of euro area differences in wealth inequality. Note that the share contributed by the income process is conditional on the order in which the country-specific parameters are set to those of

the EA benchmark country. This is due to an interdependence of all policies with each other and with labor market dynamics. To exemplify the interaction of labor market dynamics with the unemployment insurance or public pension scheme, one can consider two countries with different degrees of net earnings risk. The same net unemployment or pension replacement rate would lead to a larger crowding-out of savings in the country with low income risk, because future benefits are less uncertain. Next in column (3), I also set the unemployment insurance system of each country equal to the one in the fictive EA reference country, but each country retains its own public and occupational pension scheme and minimum income support program. Taking the difference between columns (2) and (3) reveals that the unemployment insurance system contributes 4.7%, conditional on the income process being equal across all countries. Continuing in this manner, I am ultimately able to separate the conditional role of each institution for euro area wealth inequality differences. Finally, setting all the parameters to the one of the euro area reference country will lead to an  $R$  of zero. This is because  $R$  assesses the model's forecasts,  $\hat{\lambda}_c(\boldsymbol{\theta}_c)$ , against a simple benchmark model predicting the mean,  $\bar{\lambda}$ , for each country and hence zero variation in cross-country differences in wealth inequality, a prediction that my model exactly makes if there are no euro area differences in country-specific parameters,  $\boldsymbol{\theta}_c$ .

To get an estimate of the effects of income risk and the welfare policies that is not dependent on the specific ordering, I will determine the contribution to the overall fraction explained by each factor for every possible ordering and take the average. In total, there are four different factors considered in the decomposition, which leads to 16 possible orders, each providing an estimate for the contribution of one determinant. Table 6 provides the results of this final unconditional decomposition method.

The decomposition results indeed change slightly compared to Table 5 and suggest that welfare policies contribute 57.5% to the euro area differences in the net wealth Gini coefficients for the bottom 95% of the wealth distribution. It turns out that the most important drivers of the social security system for determining wealth inequality differences across the euro area are means-tested minimum income support programs and pension schemes. While the pension system can rationalize 10.7%, minimum income support programs stand out by far, accounting for 44.8% of the differences. Institutional differences in unemployment insurance systems across the euro area, by contrast, play with 2% only a minor role. Furthermore, 12.6% of the cross-country differences in wealth inequality can be attributed to the net earnings process and unemployment dynamics.

The strong effect of minimum income support programs on the wealth distribution relative to other policies is due to several distinct features. First, means-tested minimum

Table 5: Conditional decomposition of cross-country differences in wealth inequality

	(1)	(2)	(3)	(4)	Fraction explained
MI*	-	-	-	-	42.6%
Pensions	-	-	-	set to EA	10.6%
UI*	-	-	set to EA	set to EA	4.7%
Income	-	set to EA	set to EA	set to EA	12.2%
Austria	0.634	0.600	0.601	0.586	
Belgium	0.535	0.566	0.568	0.574	
Finland	0.608	0.605	0.601	0.601	
France	0.603	0.596	0.594	0.581	
Germany	0.624	0.603	0.602	0.598	
Greece	0.525	0.507	0.519	0.528	
Italy	0.518	0.521	0.530	0.525	
Netherlands	0.607	0.629	0.620	0.619	
Portugal	0.547	0.563	0.559	0.570	
Spain	0.524	0.534	0.531	0.543	
$R$	70.1%	57.9%	53.1%	42.6%	
$\beta$	97.3	97.5	97.6	97.6	

Notes: UI\*= Unemployment insurance, MI\*=Minimum-income support.

Table 6: Unconditional decomposition of cross-country differences in wealth inequality

	Fraction explained
MI*	44.8%
Pensions	10.7%
UI*	2.0%
Income	12.6%

*Notes:* UI\* = Unemployment insurance, MI\* = Minimum-income support.

income benefits do not depend on any past contributions, but only on the households' current means. Hence, they are much more redistributive across individuals compared to unemployment benefits or pensions, which are in the euro area rather redistributive over the life cycle. Hubbard et al. (1995) have shown that the lower bound on consumption leaves the precautionary savings decision of households with high expected life-time income relatively unaffected, while the need for self-insurance for households in the lower part of the income distribution strongly reduces, thereby increasing wealth inequality. Second, minimum income assistance guarantees a certain lump sum transfer, while future potential unemployment benefits replace a constant fraction of previous net income and are hence still dependent on uncertain net income. While the receipt of pension payments during retirement is certain, also some uncertainty about the exact pension level during retirement remains, as pensions will depend on the household's pre-retirement labor market performance. Since households are risk averse, this uncertainty characteristic of unemployment and pension benefits leads to weaker effects on wealth inequality despite sizeable aggregate effects on wealth. Third, the asset-test of minimum income support introduces an implicit tax on savings, such that low-wealth households face a trade-off between saving for bad income states and dissaving to become eligible for income support. And last, minimum income benefits are of unlimited duration, while unemployment benefits mitigate earnings losses only temporarily.

It is shown in Table 7 that the importance of minimum income support programs for determining wealth inequality differences does not crucially depend on the transfers' characteristic of being asset-tested. It can be shown that the explanatory power of the model remains high at 70.8% when assuming the extreme case of 100% asset exemption levels for all euro area countries. While e.g. the predicted wealth inequality for Germany

Table 7: Robustness

Specification	R
1) No asset test	70.8%
2) Private transfers	68.2%
3) Share held by bottom 50%	69.7%

slightly worsens after abolishing the asset-test and recalibrating the time preference parameter, the model prediction of the Gini wealth coefficient for Greece moves under this assumption closer to the actual value. Next, I also relax the assumption of zero income during unemployment and allow for positive private transfers of 10% of the country's median net earnings to unemployed households. Relative to the baseline results, the explanatory power of the model only slightly decreases by 2%. As a further robustness test, the share of total net wealth held by the bottom 50% of the population is considered as an alternative measure for wealth inequality in Table 7, again excluding the top 5th percentile of the wealth distribution in the data. The results are also robust to this measure. On average, policies and income processes can also account for 69.7% of the cross-country variation in the net wealth share held by the poorest 50%.<sup>17</sup>

These results also shed light on the documented empirical puzzle that countries with a larger reduction in the income Gini coefficient through transfers, show higher wealth inequality. Since higher after-tax earnings inequality and unemployment rates lead to higher wealth inequality, one would expect that transfers by reducing income differences, lower wealth inequality in turn. In fact, the opposite is true. While transfers temporarily mitigate income differences across households, their general availability leads to a more unequal wealth distribution in the long run. Furthermore, the analysis also questions a common practice in the incomplete markets literature of estimating the standard deviation of household income shocks directly from after-tax and transfers income data for the model's calibration, as this leads to an overly low implied wealth inequality. It is therefore important to model both the earnings process and transfers explicitly if the

<sup>17</sup>While the 90/10 and the 50/10 percentile ratios are suitable and commonly used to analyze inequality of income distributions, they are less suited when working with wealth distributions, because the 10th percentile can be negative or close to zero and hence the ratio can be negative or infinite.

wealth distribution is essential for the question at hand.

## 7 Conclusion

Consistent with the theory of Hubbard et al. (1995) that public insurance distorts private savings decisions especially of low-income and low-wealth households, I empirically document that countries with more generous and redistributive welfare policies have considerably higher wealth inequality. Using a structural life-cycle model featuring labor market risk, bequests, and various institutions of the social security system, it is shown that labor market dynamics and redistributive policies can account for 70.1% of the euro area variation in wealth inequality for the bottom 95% of the wealth distribution. Furthermore, I provide a decomposition of the individual roles of the factors considered. It is shown that welfare policies can account for 57.5% of the euro area differences in wealth inequality and that the most important institution of the social security system driving these results is means-tested minimum income support provided by the government. Since minimum income support benefits are highly redistributive across individuals, certain, asset-tested and of unlimited duration, they strongly affect wealth inequality, and euro area differences in this institution can account for 44.8% of the differences in the net wealth Gini coefficients. Public and occupational pension entitlements and labor market dynamics can rationalize 10.7% and 12.6%, respectively. In contrast, cross-country differences in unemployment insurance systems have with 2% only little explanatory power. I also demonstrate that the asset-based means-testing of minimum income provision is not central to the overall explanatory power of minimum income benefits, as the model's explanatory power remains unchanged when minimum income benefits are assumed to be 100% exempt from the asset-test.

While many studies on wealth inequality focus on determinants which influence the upper tail of the wealth distribution, this analysis sheds light on the remaining part, and in particular, the role of public insurance. When considering the bottom 95% of the wealth distribution, there are still large cross-country differences in wealth inequality to be understood and this analysis reveals that redistributive welfare policies are indeed central in determining wealth inequality patterns across the euro area.

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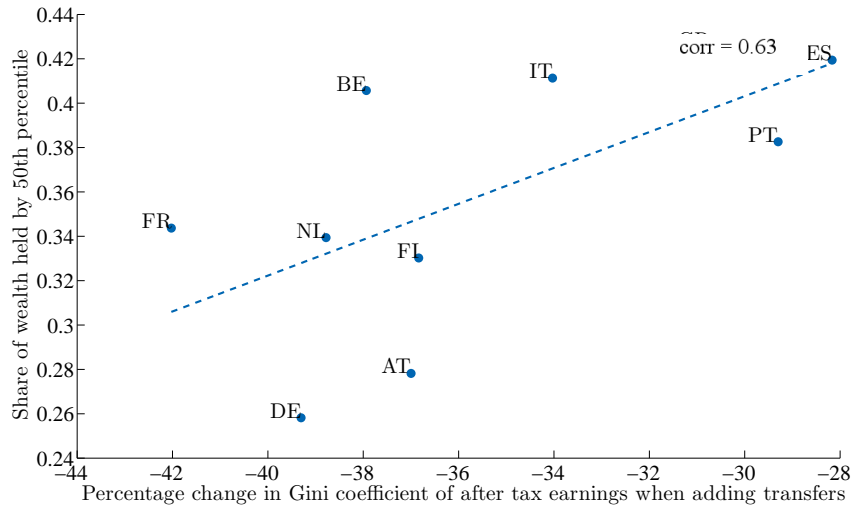
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# Appendices

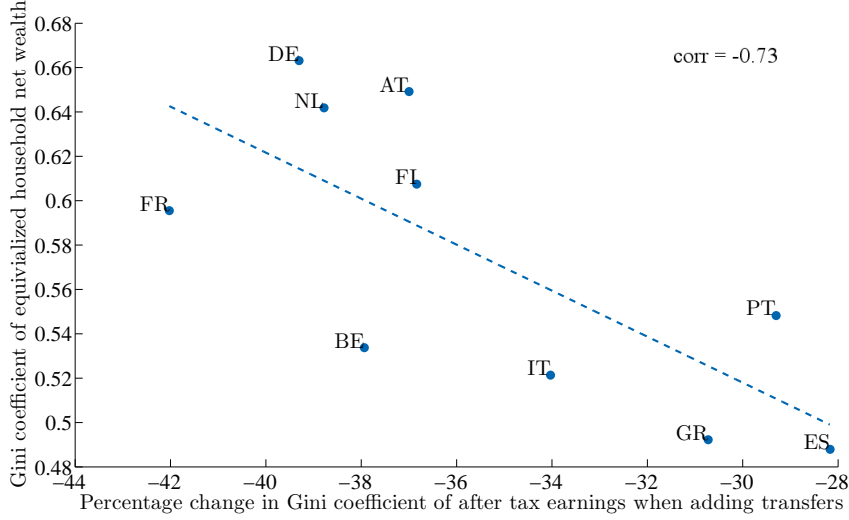
## 8 Robustness of Empirical Facts

Figure 5: Share of wealth held by 50th percentile (bottom 95%)



Sources: HFCS (~2010), EEF (2011), EU-SILC (2004-2010)

Figure 6: Gini coefficient of equivalized household net wealth (bottom 95%)



Sources: HFCS (~2010), EEF (2011), EU-SILC (2004-2010),

## 9 Numerical Methods

The household problem is solved backwards by starting in the last period of life,  $h_{62}$ . Optimal consumption and savings choices in previous periods are then derived, given subsequent optimal consumption choices and corresponding value functions. I solve for the optimal policies of households whose income is sufficiently low to be eligible for minimum income support, but whose current wealth is such that they never want to take up this support and hence choose end of period wealth holdings  $k_h > TR_c - w_h^{net}$ , by applying the endogenous gridpoint method as originally developed in ?. For lower current wealth holdings, multiple local maxima can emerge and hence, following Wellschmied (2015), I solve for the global maximum via value function iteration and allow for at least 2000 asset choices with a very fine asset grid at the low end of the asset distribution. Increasing the number of gridpoints did not have a noticeable effect on the model-implied Gini coefficients of private net wealth. Also if households are not currently eligible for minimum income support since  $w_h^{net} \geq TR_c$ , multiple maxima and distortions can arise for households with sufficiently low current wealth levels and hence optimal policies are in this region determined by value function iteration. These distortions arise due to

the life-cycle dimension and stochastic nature of earnings. Households place a positive probability on entering a low income state that could potentially make them eligible for means-tested income support in the future. Therefore, today's value function inherits the kinks in the expected value functions of states when households are eligible for minimum income benefits.

I approximate the idiosyncratic gross earnings process using a discrete Markov chain with 7 states, using the method proposed by ?.

## 10 Calibration

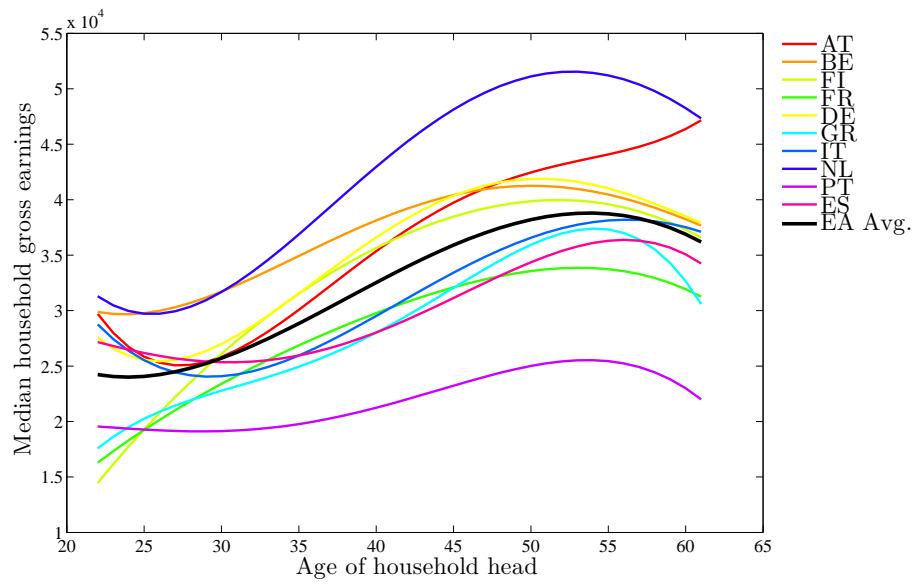
Table 8: Aggregate wealth and its distribution at the end of the life cycle

Country	$\frac{A_{mean,h \geq 84}}{A_{mean,h < 84}}$	$\frac{A_{median,h \geq 84}}{A_{median,h < 84}}$
AT	0.54	0.43
BE	0.33	0.14
FI	0.66	0.37
FR	0.50	0.27
DE	0.71	0.64
GR	0.67	0.29
IT	0.39	0.17
NL	0.59	0.36
PT	0.81	0.39
ES	0.37	0.16
EA Avg.	0.56	0.32

*Sources:* Own calculations, HFCS (~2010, excl. top 5th percentile)

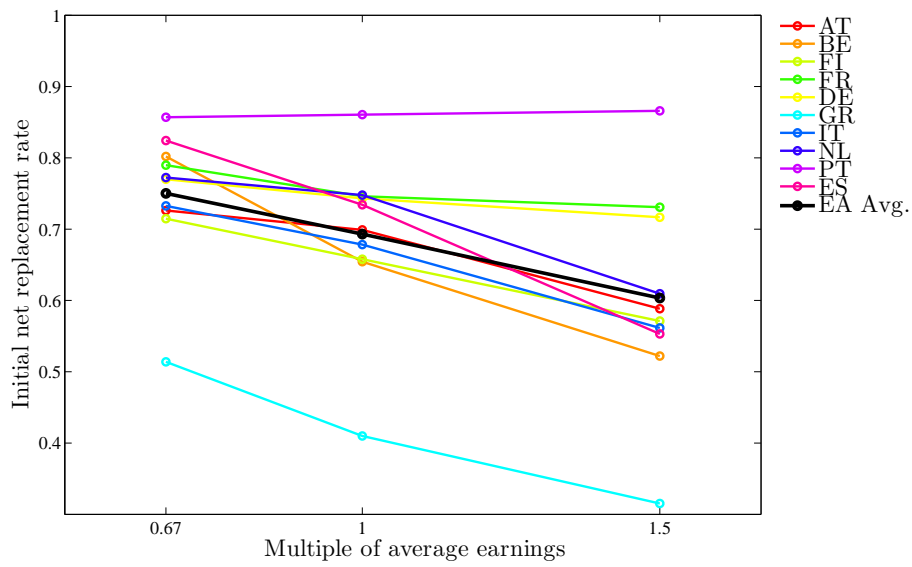
*Notes:*  $\frac{A_{mean,h \geq 84}}{A_{mean,h < 84}}$ =ratio of mean wealth of households older or equal to 84 years relative to mean wealth of households younger than 84;  $\frac{A_{median,h \geq 84}}{A_{median,h < 84}}$ =ratio of mean wealth of households older or equal to 84 years relative to mean wealth of households younger than 84. The average values for the euro area shown in the last row are used for the calibration of the bequest function.

Figure 7: Cross-sectional age-earnings profile, adjusted for euro area real wage growth and PPP



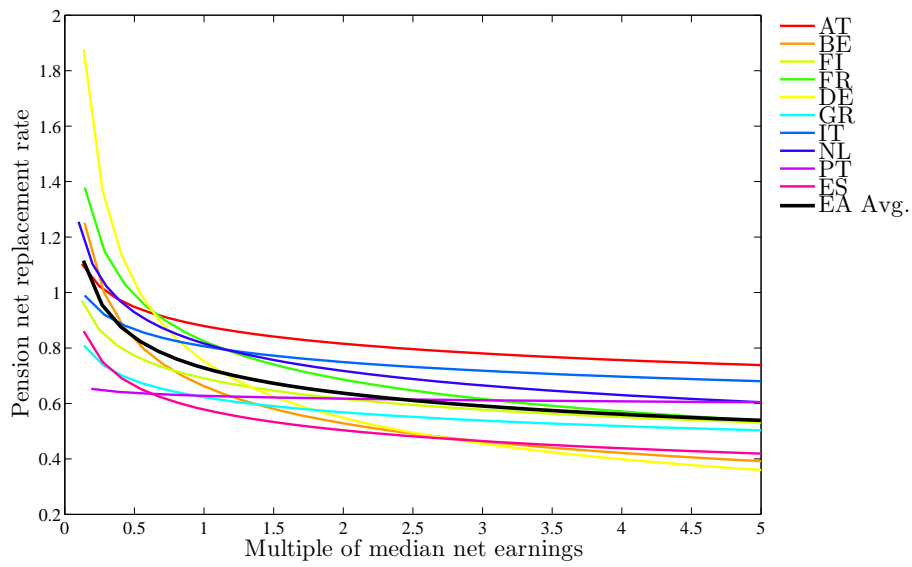
Sources: Own estimations, EU-SILC (2004-2010)

Figure 8: Unemployment net replacement rate



Sources: OECD "Benefits and Wages" database (2004/2010)

Figure 9: Public and occupational pension replacement rate



Sources: Own calculations, EU-SILC (2004-2010)