Does Risk Shape Economies? Income Volatility and Structural Change*

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Abstract

We investigate the role of risk in shaping the pattern of structural transformation of an economy. We first show a novel theoretical result: in a simple two-sector model, for any given level of GDP, higher microeconomic risk (e.g. volatility of TFP shocks) implies a smaller share of services in consumption. This occurs because higher income risk induces the representative household to increase precautionary savings, thus reducing consumption expenditure, whose level determines the structure of consumption due to non-homotheticity. The value added share of services also declines with higher risk, due to the increase in goods intensive investment relative to services intensive consumption. Time-series and cross-sectional U.S. data confirm a negative and statistically significant relationship between different measures of risk and the share of services, in both value added and consumption data. This relationship also holds in South-American and Asian countries experiencing premature de-industrialization. As these countries faced lower risk relative to the U.S. during their development, the proposed mechanism can account for part of their premature de-industrialization. Our estimates suggest that, had the U.S. experienced the same GDP volatility before WWII as it did after, its average services share would be 0.023 percentage points higher - explaining roughly 30% of the gap with premature de-industrializers at comparable income levels.

JEL Classification: L16, E21, E30.

Keywords: Strucural Change, Income Volatility, Precautionary Savings.

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1 Introduction

The productive structure of an economy affects the amount of risk that the economy displays (Acemoglu, Carvalho, Ozdaglar, and Tahbaz-Salehi, 2012, Moro, 2012, Carvalho and Gabaix, 2013, Moro, 2015). This is because volatility at the sector level transmits to the aggregate economy depending on intersectoral linkages and the relative size of sectors in aggregate consumption and in GDP. Thus, when the structure of the economy changes, even if individual sectors' volatility does not, the level of risk of the economy also changes. However, can the opposite direction of causality also emerge? If this is the case the process of structural transformation itself might be affected by the amount of risk an economy faces along the development path.

In this paper we show that indeed, in a stochastic and otherwise standard structural transformation model, the level of risk can affect the structure of the economy at a given level of GDP. The mechanism works through the interaction of precautionary savings due to income risk and non-homothetic preferences. For a given level of current income, the household makes a saving decision based on expected risk in the future. Risk enters the model through TFP shocks in the two sectors in the economy, manufacturing and services. The higher the level of microeconomic risk (i.e. the level of volatility of TFP shocks in the two sectors), the larger is income risk, and the larger is the amount of savings of the household due to precautionary motives. In a closed economy, larger savings for a certain GDP level imply smaller consumption expenditure. In turn, with non-homothetic preferences the structure of consumption depends on consumption expenditure. Thus, at any GDP level, larger risk implies larger precautionary savings, smaller consumption expenditure and, in a two-sector model with manufacturing and services, a smaller share of services. In this way, income risk shapes the structure of an economy and affects the process of structural transformation.

We then investigate the empirical relevance of the phenomenon by focusing on the U.S. and using historical data constructed in Herrendorf, Rogerson, and Valentinyi (2014). We regress the share of services in year t againsts typical controls like the level of GDP per-capita and time fixed effects, adding as a regressor the volatility of GDP. This is computed, for each t, as the standard deviation of percentage deviations from an Hodrick-Prescott filter using data of 15 years before t. Importantly, as higher GDP volatility might be simply driven by a larger manufacturing sector, which displays more volatility than services, we also control for the manufacturing share with a five years lag. Our regression specification allows to investigate whether, controlling for the level of GDP, the lagged share of manufacturing, and time fixed effects, a larger level of risk has an effect on the structure of the economy, summarized by the share of services. We consider both value added shares and consumption shares in our estimations. Across different speficifications of the model, and for both types of services shares, we find a negative and significant relationship between past volatility of GDP and current share of services, which supports the idea that risk has an effect on structural composition.

While we control for the lagged value of the manufacturing share, measured GDP volatility might still be correlated with the size of the manufacturing sector. If this is the case, the negative coefficient on volatility in our regressions would be influenced by a large manufacturing sector, which correlates with a smaller services sector.[1] For this reason, we also investigate the relationship of the share of services with another measures of aggregate risk: the Geopolitical Risk (GPR) index by Caldara and Iacoviello (2022). This index captures a dimension of expected risk faced by households, that is not directly related to the structure of the economy. It is constructed from text-based analysis of major international newspapers, quantifying the frequency of articles that discuss geopolitical tensions, such as wars, military threats, and terrorist acts. Even with this measure, the share of services correlates negatively and in statistically significant way.

We then relate the above theoretical finding to the recent literature on *premature de-industrialization*. In a nutshell, this phenomenon indicates that, at a given level of GDP per capita, countries undergoing structural transformation in more recent years tend to have a smaller share of manufacturing value added than those that transformed earlier in time. That is, in countries facing premature de-industrialization, the manufacturing share increases less and begins to decline at lower income levels compared to countries that industrialized - and subsequently deindustrialized - earlier. Different explanations have been proposed to account for premature de-industrialization. Huneeus and Rogerson (2024) show that countries experiencing premature de-industrialization may have relatively slower agricultural productivity growth profiles. Rodrik (2016) instead, argues that trade and globalization likely played a more significant role in the de-industrialization of developing countries compared to technological progress, which is typically considered a primary driver in advanced economies. Sposi, Yi, and Zhang (2024) also emphasize the role of sectoral trade integration, which, alongside sector-biased productivity growth, contribute to account for de-industrialization.

In this paper, we focus on an alternative potential source of premature de-industrialization, given by the different amount of macroeconomic risk faced during the development process by late and early industrializers. Macroeconomic risk is high for early industrializers (last part of the 19th and first half of the 20th century) while it is low for late ones (post WWII). For instance, Barro and Ursua (2008) report an average volatility of GDP for a set of OECD and

[1]Note that the two shares do not sum to one, because of the agricultural sector. However, given the small size of the agricultural sector, the two shares might be highly negatively correlated.

non-OECD countries of 6.26% for the period 1870-1947, compared to 3.36% for the period 1948-2006.[2] In the early period (1870-1947), OECD countries display a volatility of 5.95%. Non-OECD countries, which typically lag OECD ones in terms of development measures, including structural transformation, display a volatility of GDP for the post-WWI period of 4.36%. This suggests that countries that develop later (after WWII) face lower aggregate risk, compared to those that did it earlier.

We then study the pattern of the share of services in GDP as income grows for South-American and Asian coutries, which are late (in time) industrializers and premature (in income) de-industrializers (Rodrik, 2016), and compare it with the U.S. experience.[3] To do this, we run a typical panel regression with country fixed effect of the share of services as the dependent variable against the log of GDP and the log of GDP squared. In the top panel of Figure 1 we report the data (colored dots) after removing the estimated fixed effects, and the estimated relationship.[4] The share of services of the U.S. lies well below the estimated relationship. By using historical time series from Herrendorf, Rogerson, and Valentinyi (2014), the yellow triangles in the top panel of Figure 1 confirm that the share of services in the U.S. is below the estimated relationship even for lower income levels.[5] This confirms, from the perspective of the share of services, the phenomenon of premature de-industrialization pointed out in previous literature.

The bottom panel of Figure 1 then compares macroeconomic risk for premature de-industrializers with the U.S. The black line is the estimated relationship of a panel regression with country fixed effects between GDP volatility as the dependent variable against the log of GDP and the log of GDP squared, where the data coverage is the same as for the top panel of Figure 1. GDP volatility in year t for country j is computed as the standard deviation of percentage deviations of GDP from an Hodrick-Prescott in the 15 years before t. The colored dots report the data used for the estimation after removing country fixed effects. As well documented in Acemoglu and Zilibotti (1997), Koren and Tenreyro (2007) and Moro (2015), the bottom panel of Figure 1 confirms that volatility declines as GDP and the share of services grow. The yellow triangles in the figure report historical U.S. volatility.[6]

[2]See their Table 3, last column. We compute the 1870-1947 period from their reported values for the 1870–1913 and the 1914–1947.

[3]We choose to use the share of services instead of the share of manufacturing as its monotonic behavior makes it easier to visualize premature de-instrialization, measured here as a larger share of services with respect to the U.S. at each GDP level. For evidence of premature de-industrialization using the share of manufacturing see Rodrik (2016), Huneeus and Rogerson (2024) and Sposi, Yi, and Zhang (2024).

[4]We use data from the GGDC 10-Sector Database and Penn World Table version 10.01 from 1950 to 2010 for all countries in Figure 1.

[5]The historical U.S. share in Figure 1 (yellow triangles) is normalized such that it coincides with modern data (green dots) in the year 1999.

[6]As for the top panel of Figure 1, the historical data for the U.S. (yellow triangles) are normalized such

It stands out that the volatility of GDP experienced by the U.S. has been larger than that of late industrializers at many stages of development. This suggests that the theoretical mechanism proposed in this paper might have empirical relevance in accounting for part of premature de-industrialization of South-American and Asian countries with respect to the U.S. experience.

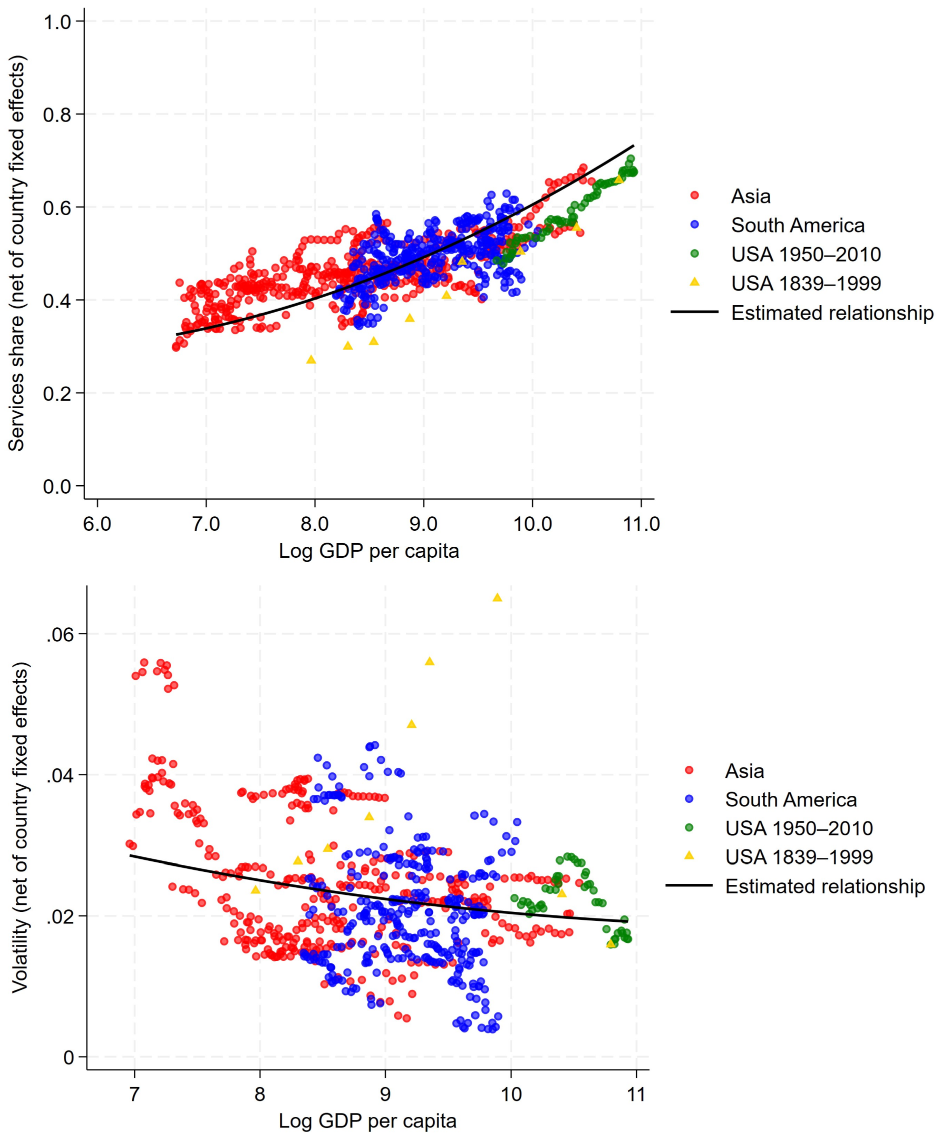

Figure 1: Top panel: Share of Services in GDP and GDP per-capita. Bottom panel: GDP volatility and GDP per-capita. Asia, South-America and U.S. 1950 data are from the GGDC 10-Sector Database and Penn World Table version 10.01. U.S. 1839-1999 are from Herrendorf, Rogerson, and Valentinyi (2014).

that they coincides with modern data (green dots) in the year 1999.

Motivated by the evidence in Figure 1, we investigate empirically whether risk has a role in shaping the structure of the economy of premature deindustrializers in South-America and Asia. We exploit countries heterogeneity, running panel regressions that control for both country and time fixed effects. We find that for the group of countries as a whole, there is a negative and statistically significant relationship between the share of services at time t and GDP volatility of the same country computed using the 15 years before t. The relationship is maintained when considering only the group of South-American countries. When considering the group of Asian countries alone, the relationship is still negative, but not statistically significant. This is due to the heterogeneity in the group, that contains economies with high intensity in manufacturing exports, which are typically insulated from premature de-industrialization (Rodrik, 2016) and countries with a low intensity in manufacturing exports. By splitting the sample according to high versus low manufacturing exports intensity, we find that the relationship is maintained and is statistically significant for the latter group, but not for the former. This confirms that our mechanism is more likely to have empirical relevance in economies where the manufacturing share is less driven by exports.

In the last section of the paper we report GDP volatility of premature de-industrializers and the U.S., showing that post-WWII, volatility in South American and Asian countries is similar to that of the U.S. in the same period, but substantially lower than U.S. pre-war levels, when the share of manufacturing in GDP in the U.S. was expanding. In light of our theory, this observation suggests that the U.S. can be considered a *late* (in income) *industrializer* due to its high volatility during the industrialization period. In a counterfactual exercise we ask what would have been structural transformation in the U.S. had this country displayed the same GDP volatility in the pre- as in the post-WWII period. Based on our time-series estimates, the counterfactual suggests the share of services would have been on average 0.023 larger. This represents about 31% of the observed difference in the share of services between the U.S. and the group of premature de-industrializers at similar income levels.

The reminder of the paper is as follows. In section 2, we discuss the related literature; in section 3 we present a simple two-sector-two-period stochastic model that allows to show the relationship between risk and structural composition; in section 4 we present the empirical evidence on the U.S. and in section 5 the cross-country evidence. In section 6 we use our results to measure how much of the difference in the share of services between the U.S. and premature de-industrializers can be accounted for by our mechanism. Finally, in section 7 we conclude.

2 Related Work

There is a large literature on the determinants of structural transformation.[7] These can be grouped broadly in four main groups: income effects, which suggest that, as income grows, the relative demand of services to manufactured goods increase (Kongsamut, Rebelo, and Xie, 2001, Boppart, 2014, Comin, Lashkari, and Mestieri, 2021); substitution effects coupled with faster technological change in manufacturing relative to services, and in agriculture relative to manufacturing and services (Baumol, 1967, Ngai and Pissarides, 2007,); marketization, which implies that services previously produced at home become purchased in the market (Ngai and Pissarides, 2008, Ngai and Petrongolo, 2017, Moro, Moslehi, and Tanaka, 2017); and openess of an economy (Uy, Yi, and Zhang, 2013).

A subset of this literature studied how the structural composition of an economy and its evolution over time (i.e. structural transformation) affects aggregate risk displayed by that economy (Acemoglu, Carvalho, Ozdaglar, and Tahbaz-Salehi, 2012, Moro, 2012, Carvalho and Gabaix, 2013, Moro, 2015, Rubini and Moro, 2024). To the best of our knowledge, however, our paper is the first to investigate the causal link going from risk to the structure of the economy, thus proposing a new theoretical mechanism which determines structural transformation. In doing so, we relate mainly to both the above literature on structural change and that on precautionary savings.

Precautionary saving refers to the extra wealth that individuals accumulate in response to uncertainty about future income. In the classic life-cycle or permanent-income framework, an increase in uncertainty (e.g. riskier future income) leads forward-looking consumers to cut current consumption and raise saving as self-insurance (Kimball, 1990). In our setting, these changes in consumption trigger changes in the structure of the economy, due to non-homothetic preferences and a different composition of consumption and investment.

Precautionary savings have been estimated in different contexts, and their quantitative relevance appear to vary substantially with income, institutional context and cultural factors. In studies for the U.S. that employ relatively recent datasets, estimates of precautionary savings are in the single-digit percentages for the general population (Parker and Preston, 2005). For the U.K., instead, Banks, Blundell, and Brugiavini (2001) provide strong evidence for a precautionary saving behavior driven by income risk, with substantial effects on consumption growth. In Germany, the reunification provided a test of precautionary behavior: East Germans, after 1990, faced a new uncertain environment and indeed increased their savings markedly compared to West Germans. Fuchs-Schündeln (2008) calibrates a life-cycle model to match East- and West-German saving patterns, finding that a precautionary saving mo-

[7] See Herrendorf, Rogerson, and Valentinyi (2014) and Moro and Valdes (2021) for surveys of the literature.

tive is key to explain the higher saving of East Germans and the convergence of their saving rates toward West German levels over the 1990s.

More dramatic examples of precautionary savings in recent decades are found in China, which has witnessed persistently high household saving rates (30–40% of disposable income) since the 1990s. In He, Huang, Liu, and Zhu (2018), a large-scale reform of state-owned enterprises (SOEs) in the late 1990s – which led to millions of layoffs and ended the "iron rice bowl" job security – is used as a natural experiment. Households associated with the reformed SOEs suddenly faced higher income uncertainty. As a result, their saving rates jumped. Precautionary savings accounted for about 40% of the wealth accumulated by urban SOE-affiliated households between 1995 and 2002. This result suggests that when income risk is high in middle-income countries, precautionary savings can be substantially larger than levels observed in the U.S.

Considering directly the relationship between aggregate risk and savings, Norman, Schmidt-Hebbel, and Serven (2000) find that indicators of macroeconomic volatility or uncertainty are positively correlated with national saving rates. Recently, Georgarakos, Kim, Coibion, Shim, Lee, Gorodnichenko, Kenny, Han, and Weber (2025) report that the willingness to pay to remove business cycles (i.e. aggregate risk) is substantially larger than that calculated in Lucas (2003) and heterogenous across economies: respondents in crisis-scarred countries like Korea and Greece report an average willingness to pay to avoid business cycles of around 8% and 6% of lifetime consumption, respectively, whereas those in Belgium and the Netherlands are only willing to sacrifice 3-4%.

Thus, the precautionary saving motive appears present but also heterogeneous across countries. In rich countries, it plays a measurable but quantitatively small role in wealth accumulation. In countries or subsets of the population facing high risk – whether due to economic transition, weak insurance, or macro volatility – precautionary saving can be very large, accounting for a substantial share of total savings.

3 Model

3.1 Environment

We study a two-sector (manufacturing and services) and two-period (0 and 1) structural change model in a stochastic environment. Uncertainty enters the model through stochastic total factor productivity (TFP). In period 0 there is no uncertainty and all the fundamentals of the economy are known, including the probability distribution of TFP shocks that will be realized in period 1. There are four possible states associated to the realizations of the two TFP terms in period 1, which are extracted from two binary distributions, potentially correlated, and can take values $A_{m,1} \in \{a - \varepsilon_m, a + \varepsilon_m\}$ in the manufacturing sector and $A_{s,1} \in \{a - \varepsilon_s, a + \varepsilon_s\}$ in the services sector, where a, ε_m, and ε_s are positive parameters, representing the common mean and the two standard deviations of the two distributions. The probabilties of occurrence of the four states are given by $P_1 = P(a + \varepsilon_m \cap a + \varepsilon_s)$, $P_2 = P(a + \varepsilon_m \cap a - \varepsilon_s)$, $P_3 = P(a - \varepsilon_m \cap a + \varepsilon_s)$, and $P_4 = P(a - \varepsilon_m \cap a - \varepsilon_s)$.

3.2 Firms

There is a price-taking firm acting in perfect competition in each sector, which uses capital to produce output using the following technology:

$$y_{i,t} = A_{i,t} k_{i,t}, \;\; i = m, s, \tag{1}$$

where A_i is the technological level of the firm and k_i is the amount of capital used in production.[8] In each period, the firm in each sector solves the following profit-maximization problem

$$\underset{k_{i,t}}{max}\, \pi_{i,t} = p_{i,t} \left(A_{i,t} k_{i,t} \right) - r_t k_{i,t}. \tag{2}$$

From the first order condition of the firm's problem we have that

$$p_{i,t} A_{i,t} = r_t,$$

so that

$$\frac{p_{m,t}}{p_{s,t}} = \frac{A_{s,t}}{A_{m,t}}. \tag{3}$$

Total output (i.e. GDP) in the economy in manufacturing units can be expressed as either side of the following expression

$$y_{m,t} + \frac{p_{s,t}}{p_{m,t}} y_{s,t} = \frac{r_t}{p_{m,t}} k_t, \tag{4}$$

where $k_t = k_{m,t} + k_{s,t}$. Finally, substituting (1) and (3) into (4) we can write

$$A_{m,t} k_t = \frac{r_t}{p_{m,t}} k_t. \tag{5}$$

[8] Note that the firm i knows her technology level $A_{i,t}$ when solving her maximization problem. This implies that in period 1 the firm maximizes after the realization of the technological shock.

3.3 Household

There is a representative household living for two periods $t = 0, 1$ who has istantaneous preferences over manufacturing $c_{m,t}$ and services $c_{s,t}$ given by the Stone-Geary aggregator

$$c_t = [\omega_m^{1/\varepsilon} c_{m,t}^{\frac{\varepsilon-1}{\varepsilon}} + \omega_s^{1/\varepsilon}(c_{s,t}+s)^{\frac{\varepsilon-1}{\varepsilon}}]^{\frac{\varepsilon}{\varepsilon-1}},$$

where s is a positive parameter. Each period the household owns an amount of capital k_t that is rented to firms in the market in exchange for the rental rate r_t. The household starts period 0 with an amount of capital k_0 and can buy manufacturing goods to build capital for period 1, k_1. There is full depreciation so capital k_0 disappears after production in period 0. Denote q as the state of the world in period 1, which is uncertain in period 0. The household then solves

$$\max_{\{c_{m,0},c_{s,0},k_1,c_{m,1}(q),c_{s,1}(q)\}} \left\{ \frac{c_0^{1-\sigma}}{1-\sigma} + \beta E_0 \left[\frac{c_1(q)^{1-\sigma}}{1-\sigma} \right] \right\}$$

subject to

$$c_t = [\omega_m^{1/\varepsilon} c_{m,t}^{\frac{\varepsilon-1}{\varepsilon}} + \omega_s^{1/\varepsilon}(c_{s,t}+s)^{\frac{\varepsilon-1}{\varepsilon}}]^{\frac{\varepsilon}{\varepsilon-1}} \ \forall t = 0, 1,$$

$$p_{m,0}c_{m,0} + p_{s,0}c_{s,0} + p_{m,0}k_1 = r_0 k_0,$$

$$p_{m,1}(q)c_{m,1}(q) + p_{s,1}(q)c_{s,1}(q) = r_1(q)k_1 \ \forall q.$$

The household problem can be split into two parts. First, we maximize the consumption index at each t

$$\max_{c_{m,t},c_{s,t}} c_t = [\omega_m^{1/\varepsilon} c_{m,t}^{\frac{\varepsilon-1}{\varepsilon}} + \omega_s^{1/\varepsilon}(c_{s,t}+s)^{\frac{\varepsilon-1}{\varepsilon}}]^{\frac{\varepsilon}{\varepsilon-1}},$$

subject to an expenditure constraint

$$p_{m,t}c_{m,t} + p_{s,t}c_{s,t} = \bar{w},$$

where $\bar{w}$ is an exogenous expenditure level. This problem delivers as solution

$$c_{s,t} = \frac{\frac{\bar{w}}{p_{m,t}}\left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon} - \frac{\omega_m}{\omega_s}s}{\frac{\omega_m}{\omega_s} + \left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon-1}},$$

$$c_{m,t} = \left(\frac{p_{m,t}}{p_{s,t}}\right)^{-\varepsilon} \frac{\omega_m}{\omega_s}(c_{s,t}+s).$$

Also, the first order conditions of the static problem allow to show that

$$p_{m,t}c_{m,t} + p_{s,t}c_{s,t} = p_t c_t - p_{s,t}s,$$

where

$$p_t = \left[\omega_g \left(p_{m,t}\right)^{1-\varepsilon} + \omega_s \left(p_{s,t}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}},$$

is the price index of the consumption index c_t, so that we can write

$$c_{s,t} = \frac{\frac{p_t c_t - p_{s,t}s}{p_{m,t}} \left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon} - \frac{\omega_m}{\omega_s} s}{\frac{\omega_m}{\omega_s} + \left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon-1}}. \tag{6}$$

By using the above results from the static problem, we can then rewrite the original problem as

$$\underset{\{c_0, c_1(q), k_1\}}{max} \left\{\frac{c_0^{1-\sigma}}{1-\sigma} + \beta E_0 \left[\frac{c_1(q)^{1-\sigma}}{1-\sigma}\right]\right\},$$

subject to

$$p_0 c_0 + p_{m,0}k_1 = p_{m,0}A_{m,0}k_0 + p_{s,0}s,$$

$$p_1(q)c_1(q) = p_{m,1}(q)A_{m,1}(q)k_1 + p_{s,1}(q)s,$$

where we used the fact that $p_{m,0}A_{m,0}k_0 = r_0 k_0$ and $p_{m,1}(q)A_{m,1}(q)k_1 = r_1(q)k_1$ for each state of the world in period 1.

Using the first order conditions and the constraints to substitute for c_0 and $c_1(q)$, the Euler equation for the problem is

$$\left(\frac{A_{m,0}k_0 + \frac{p_{s,0}}{p_{m,0}}s - k_1}{p_0/p_{m,0}}\right)^{-\sigma} \frac{1}{p_0/p_{m,0}} = \beta E\left[\left(\frac{A_{m,1}(q)k_1 + \frac{p_{s,1}(q)}{p_{m,1}(q)}s}{p_1(q)/p_{m,1}(q)}\right)^{-\sigma} \frac{A_{m,1}(q)}{p_1(q)/p_{m,1}(q)}\right] \tag{7}$$

where

$$p_{s,0}/p_{m,0} = A_{m,0}/A_{s,0},$$

$$p_{s,1}(q)/p_{m,1}(q) = A_{m,1}(q)/A_{s,1}(q),$$

$$p_0/p_{m,0} = \left\{\omega_m + \omega_s \left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon}\right\}^{\frac{1}{1-\varepsilon}},$$

$$p_1(q)/p_{m,1}(q) = \left\{\omega_m + \omega_s \left[\frac{p_{s,1}(q)}{p_{m,1}(q)}\right]^{1-\varepsilon}\right\}^{\frac{1}{1-\varepsilon}}.$$

Equation (7) can be solved numerically for the amount of savings in period 0, k_1.

3.4 Volatility and sectoral composition

We are interested in the effect of different risk levels in the economy on the determination of the share of services in consumption in period 0, defined as

$$\frac{p_{s,0}c_{s,0}}{p_{m,0}c_{m,0} + p_{s,0}c_{s,0}}, \tag{8}$$

and in value added, defined as

$$\frac{p_{s,0}c_{s,0}}{p_{m,0}A_{m,0}k_0}. \tag{9}$$

A larger volatility faced by the household implies a larger value of k_1, due to precautionary savings. In turn, for a given income level in period 0, $p_{m,0}A_{m,0}k_0$, a larger value of k_1 implies a smaller value of consumption expenditure in period 0, as made clear by the budget constraint, conveniently re-written as follows

$$p_{m,0}c_{m,0} + p_{s,0}c_{s,0} = p_{m,0}A_{m,0}k_0 - p_{m,0}k_1.$$

As the model displays non-homothetic preferences, a smaller level of consumption expenditure implies a smaller share of services in consumption in period 0. Thus, at the same level of GDP, the level of risk in the economy determines the size of the share of services in consumption in the economy.

In addition, equation (9) shows that also the value added share of services declines as volatility increases. This is because the denominator is invariant to volatility in period 1, while the numerator declines with volatility, due to the reduction in consumption expenditure and the increase in precautionary savings.

The effects of risk on savings, consumption expenditure and the share of services in both consumption and value added are shown in Figure 2, for the parametrization reported in Table 1. The increase in risk in Figure 2 is achieved by increasing the value of both ε_m and ε_s from the initial value displayed in Table 1. We consider, however, three different specifications for the correlation of shocks. First, we consider the case of perfectly uncorrelated shocks between the two sector. This is achieved by setting the probabilities of the four states as: $P_1 = 0.25$, $P_2 = 0.25$, $P_3 = 0.25$, and $P_4 = 0.25$. The case of perfectly correlated shocks in the two sectors implies $P_1 = 0.5$, $P_2 = 0$, $P_3 = 0$, and $P_4 = 0.5$, while the case of perfect negative correlation between the shocks implies $P_1 = 0$, $P_2 = 0.5$, $P_3 = 0.5$, and $P_4 = 0$. Figure 2 shows that, for a given level of GDP in period 0, determined by initial capital k_0 and initial TFP levels $a_{m,0}$ and $a_{s,0}$, an increase in risk in the economy transmits into higher

Table 1: Models parameters

Parameters	Symbol	Value
Risk aversion	σ	3.5
Subjective Discount factor	β	0.95
Non-homothetic component in services	s	0.1
Elasticity between manufacturing and services	ε	0.5
Initial capital	k_0	1
Stone-Geary consumption weights	$\omega_{c,m} = \omega_{c,s}$	0.5
Total factor productivities (at $t = 0$)	$a_{m,0} = a_{s,0}$	2

savings in period 0 (i.e. capital available at period 1), a decline in consumption expenditure at $t = 0$, and a decline in both the share of services in consumption expenditure and in the share of services in total value added at $t = 0$. Thus, the amount of risk affects the structure of the economy through its effect on precautionary savings.

In Appendix B we generalize the result to the case in which investment is a composite of both manufacturing and services. As long as investment is more intensive in manufacturing than consumption, which is the empirically relevant case (García-Santana, Pijoan-Mas, and Villacorta, 2021), the effects are qualitatively similar to those in Figure 2. In Appendix B we also generalize the model to Epstein-Zin-Weil type of preferences, and show that a larger value of risk aversion, for a given elasticity of intertemporal substitution, increases the negative effect on the share of services, as it increases the amount of precautionary savings.

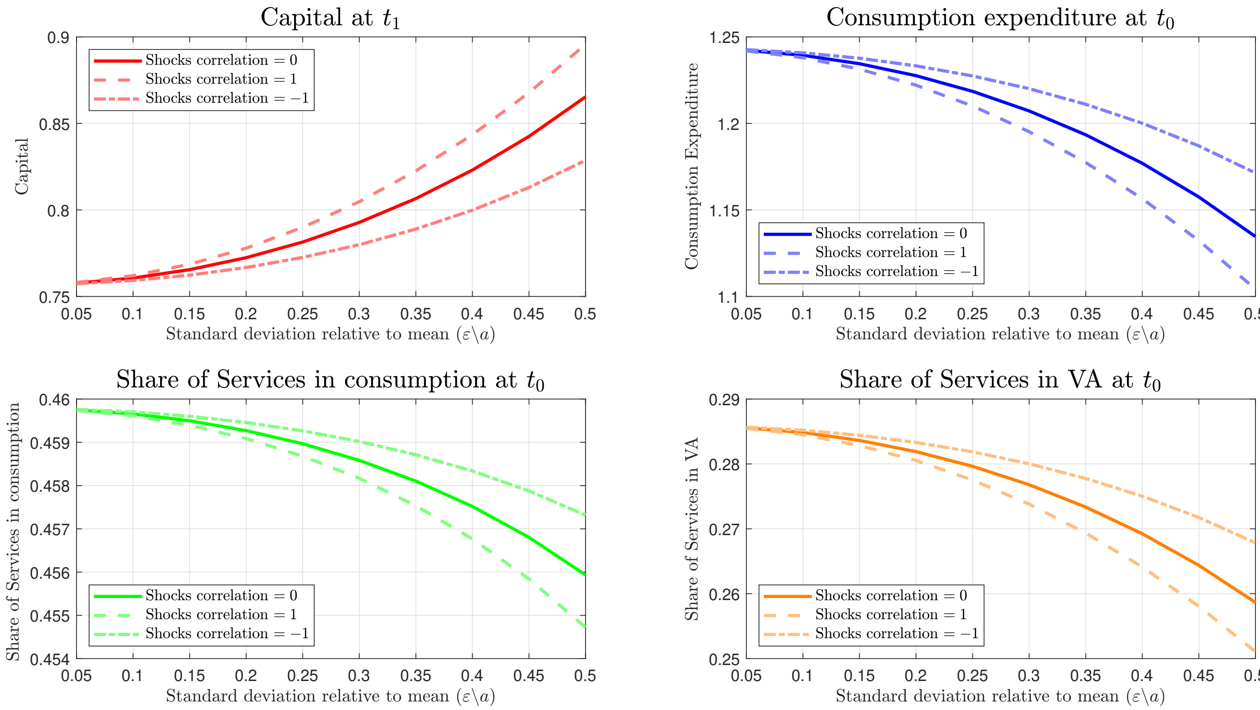

Figure 2: The effect of risk on the economy in period $t = 0$.

In Figure 3 we show that the current structure of the economy (i.e. the share of services at $t = 0$) does not affect the mechanism. The figure reports the four variables for three different TFP levels and different levels of risk. The three TFP levels are set such that TFP grows in both sectors, but at a faster pace in manufacturing (10% growth) with respect to services (5% growth). As TFP grows in both sectors, there is structural transformation due to both income and substitution effects, which is summarized in each panel as the (initial) continuous line moves upwards to reach the dashed and then the dotted line. Regardless of the current level of TFP, which determines on which of the three lines the economy finds itself, an increase in risk has the effect of reducing the share of services.

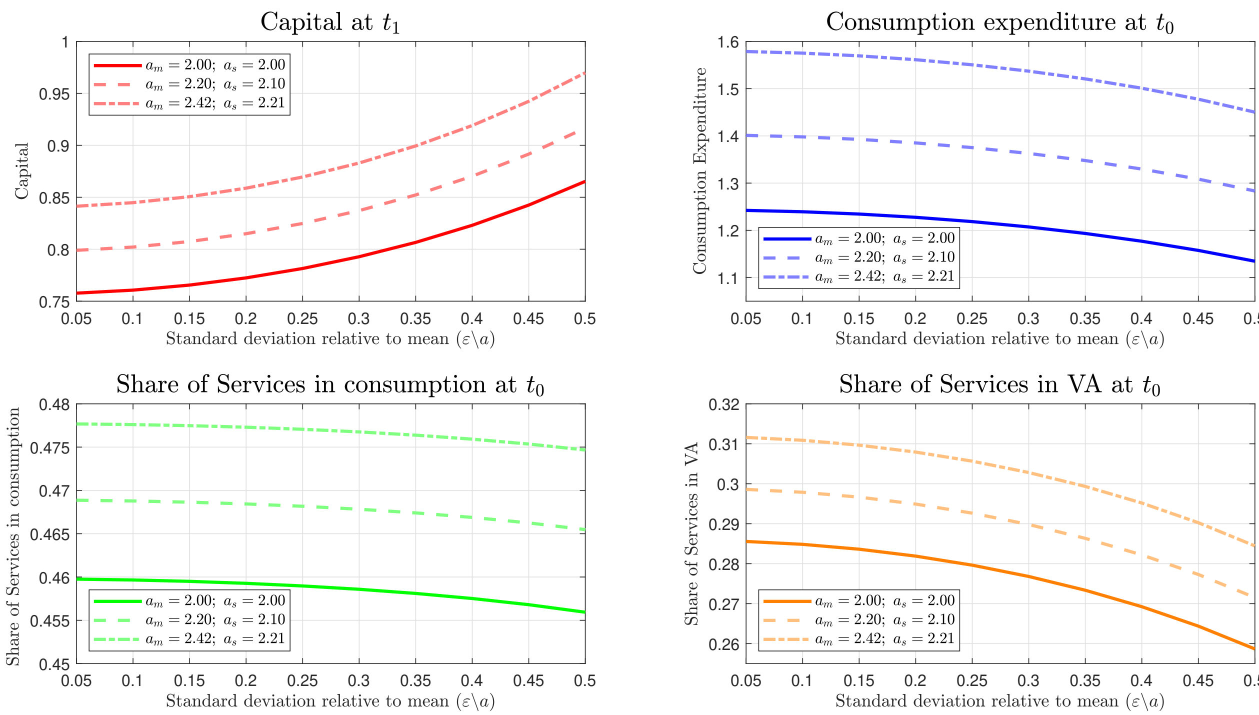

Figure 3: Structure of the economy and the effect of risk

4 U.S. Evidence

4.1 Time Series Evidence

In this section, we investigate whether income volatility, through its effect on precautionary savings, influenced the pace of structural change in the U.S. Such empirical investigation of the phenomenon for a single country requires sufficiently long time series on i) some measure of income risk faced by households in an economy, and ii) the structural composition of the economy. Also, the economy under study should ideally experience substantial structural transformation during the period of analysis. Measures of income risk at the individual level are typically hard to find, also for developed countries with leading statistical agencies like the U.S. For this reason, we focus on the measure of income risk given by GDP volatility.

For our purposes, the historical time series compiled by Herrendorf, Rogerson, and Valentinyi (2014) for several countries can be used. They reconstruct data on GDP per capita and major sectors value added shares for Belgium, Finland, France, Japan, Korea, Netherlands, Spain, Sweden, the U.K. and the U.S. We focus on the U.S. for the following reasons. Among all the countries, many do not have entire coverage on sectoral shares over the twentieth century (Belgium, France, Japan, Korea, Netherlands, Spain, U.K.). Among countries with full coverage, Finland was a small economy at the start of the twentieth century, and belonged to the Russian Empire until the Russian Revolution of 1917, when it declared its independence.[9] Sweden was also a small country in 1900, and belonged to a currency union based on the gold standard with Denmark from 1873 until World War I, when the union was suspended.[10] For these reasons, we focus on the U.S., an economy that undergone a substantial process of structural transformation during the twentieth century. At the same time, it did not have major capital disruptions due to either WWI or WWII, and did not belong to unions of countries either at a political level or at the level of the currency, all factors that might have substantially affected GDP volatility in other countries.

The U.S. allows us to construct a GDP per-capita volatility series that starts in 1909 and runs until 2008, together with consistent time series on real GDP per capita and sectoral composition for the same period. Also, both value added shares and consumption shares as a measure of structural composition of the economy are available for the U.S. This allows us to test the empirical implication of the theoretical argument in section 2: *ceteris paribus*, the higher income risk the lower share of services that the country should display.

To test our hypothesis, we estimate the following regression model:

$$ser_t = \alpha + \beta_1 ln(y_t) + \beta_2 man_{t-5} + \beta_3 \sigma_{t:-15} + \beta_4 (ln(y_t))^2 + \beta_5 (man_{t-5})^2 + \sum_T \phi_T D_T + \epsilon_t,$$

where ser_t, denotes the share of the service sector's contribution, expressed as a percentage to the total nominal GDP; $ln(y_t)$, is the log of the yearly real gross domestic product per capita; $\sigma_{t:-15}$ is the risk measure, constructed as the standard deviation of percentage deviations of y_t from an HP filter in the fifteen years before year t;[11] man_{t-5} is the lagged share of the manufacturing sector's in nominal GDP; $(ln(y_t))^2$ is the square of the log of the yearly real gross domestic product per capita; $(man_{t-5})^2$ is the square of the lagged share of the manufacturing sector's contribution in nominal GDP squared; D_T are twenty-year time dummies.

Table 2 reports the regression results across different model specifications when using value added shares. All of them include $ln(y_t)$, $\sigma_{t:-15}$ and man_{t-5} as core regressors, except models (7) and (8), which exclude the lagged manufacturing share. Note that we include the latter among regressors in specification 1-6 for the following reason. Moro (2012), Carvalho

[9] Finland's population was 2.6 millions in 1900. https://stat.fi/tup/suoluk/suoluk_vaesto_en.html

[10] Sweden had a population of 5.1 millions in 1900. https://www.scb.se/en/finding-statistics/statistics-by-subject-area/population-and-living-conditions/population-composition-and-development/population-statistics/pong/tables-and-graphs/population-statistics---summary/population-and-population-changes-17492024/

[11] See Appendix A for details on the calculation of volatility.

Table 2: U.S. 1910-2008

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$ln(y_t)$	0.0825***	0.00014	-1.424***	0.0705***	-1.155*	-1.168*	0.0084	-1.175***
	(0.0122)	(0.0250)	(0.497)	(0.0135)	(0.601)	(0.611)	(0.0257)	(0.389)
$man_{i,t-5}$	-0.417***	-0.101	0.0136	-3.203***	-1.165	-1.854*		
	(0.0884)	(0.137)	(0.172)	(0.750)	(0.941)	(1.103)		
$\sigma_{i,t:t-15}$	-0.828***	-0.966***	-0.581***	-1.124***	-0.743***	-0.895***	-0.975***	-0.698***
	(0.250)	(0.162)	(0.159)	(0.272)	(0.187)	(0.135)	(0.157)	(0.142)
$D_{1930-1949}$		0.0404**				0.0448**	0.0506***	0.0543***
		(0.0184)				(0.0184)	(0.0187)	(0.0183)
$D_{1950-1969}$		0.0641***				0.0800***	0.0679***	0.0801***
		(0.0229)				(0.0223)	(0.0237)	(0.0250)
$D_{1970-1989}$		0.104***				0.112***	0.108***	0.104***
		(0.0290)				(0.0262)	(0.0310)	(0.0279)
$D_{1990-2008}$		0.175***				0.140***	0.183***	0.139***
		(0.0397)				(0.0288)	(0.0400)	(0.0285)
$ln(y_t)^2$			0.0815***		0.0667**	0.0630*		0.0645***
			(0.0264)		(0.0321)	(0.0329)		(0.0203)
$man^2_{i,t-5}$				4.698***	1.855	3.025*		
				(1.283)	(1.388)	(1.599)		
Constant	0.0239	0.628***	6.819***	0.549**	5.793**	6.268**	0.512**	5.907***
	(0.134)	(0.219)	(2.286)	(0.218)	(2.670)	(2.701)	(0.226)	(1.846)
Observations	95	95	95	95	95	95	99	99
R-squared	0.804	0.850	0.833	0.822	0.834	0.870	0.848	0.859
Time FE	NO	YES	NO	NO	NO	YES	YES	YES

Note: * p<0.1; ** p<0.05; *** p<0.01. Standard errors are in parentheses below coefficients.

Table 3: U.S. 1910-2008. Consumption data

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$ln(y_t)$	0.111***	0.0424**	-0.651*	0.109***	-0.861*	-1.183***	0.0441**	-1.027***
	(0.0115)	(0.0162)	(0.350)	(0.0119)	(0.473)	(0.412)	(0.0172)	(0.322)
man_{t-5}	-0.539***	-0.282***	-0.321***	-1.015**	0.597	1.216**		
	(0.0700)	(0.0944)	(0.101)	(0.465)	(0.791)	(0.579)		
$\sigma_{t:t-15}$	-0.663***	-1.074***	-0.538***	-0.714***	-0.412***	-0.847***	-1.254***	-1.003***
	(0.160)	(0.162)	(0.120)	(0.176)	(0.154)	(0.162)	(0.159)	(0.151)
$D_{1930-1949}$		0.0683***				0.0738***	0.0758***	0.0791***
		(0.0140)				(0.0141)	(0.0137)	(0.0134)
$D_{1950-1969}$		0.0682***				0.0736***	0.0665***	0.0775***
		(0.0159)				(0.0161)	(0.0167)	(0.0172)
$D_{1970-1989}$		0.0983***				0.0866***	0.101***	0.0977***
		(0.0200)				(0.0184)	(0.0219)	(0.0197)
$D_{1990-2008}$		0.148***				0.118***	0.171***	0.131***
		(0.0265)				(0.0220)	(0.0273)	(0.0229)
$ln(y_t)^2$			0.0412**		0.0527**	0.0671***		0.0584***
			(0.0185)		(0.0253)	(0.0222)		(0.0172)
man^2_{t-5}				0.804	-1.444	-2.134**		
				(0.797)	(1.194)	(0.866)		
Constant	-0.362***	0.140	3.076*	-0.273*	3.876*	5.447***	0.0358	4.918***
	(0.128)	(0.142)	(1.632)	(0.161)	(2.097)	(1.829)	(0.151)	(1.492)
Observations	95	95	95	95	95	95	99	99
R-squared	0.901	0.941	0.906	0.902	0.907	0.947	0.939	0.946
Year FE	NO	YES	NO	NO	NO	YES	YES	YES

Note: * $p<0.1$; ** $p<0.05$; *** $p<0.01$. Standard errors are in parentheses below coefficients.

and Gabaix (2013) and Moro (2015) document that the manufacturing sector is more volatile than the services sector. Thus, a large manufacturing sector might induce both a large level of GDP volatility and a small share of services (simply due to the fact that the manufacturing share is large). By controlling for the share of manufacturing we rule out this possibility. Models (7) and (8) also show the results omitting the share of manufacturing from the regressors, in both its linear and quadratic form.[12]

The results in Table 2 show that, across all model specifications, the coefficient on GDP volatility is negative and statistically significant. Thus, controlling for the level of GDP, a larger volatility of GDP in the previous periods has a negative effect on the share of services, which supports the mechanism discussed in section 3. The relationship holds when

[12]Note that the share of manufacturing and that of services do not sum to one, due to the share of agriculture.

controlling for quadratic terms in GDP and lagged manufacturing share, as well as time fixed effects. The stability of the negative correlation between the service share and GDP volatility across specifications reinforces the interpretation that higher volatility negatively affects the share of services, playing a key role in shaping the structure of an economy.

Table 3 reports similar estimates as in Table 2, but now the consumption share of services is used as the dependent variable, instead of the value added share of services. Also in this case, and consistent with the predictions of the model, the negative effect of volatility on the share of services is mantained across all specifications.

4.2 Other Measures of Risk for the U.S.

To test the robustness of our findings, we complement our baseline analysis based on GDP volatility with an alternative proxy for uncertainty: the Geopolitical Risk (GPR) developed by Caldara and Iacoviello (2022). This index measures a type of "expected" risk that is not directly observable through macroeconomic GDP volatility alone. In fact, while GDP volatility is a measure of realized risk (so measured ex-post), the GPR is a measure of expected risk, which is arguably the most relevant (and model consistent) measure to determine the extent of precautionary savings by households.[13] The GPR index is constructed from text-based analysis of major international newspapers, quantifying the frequency of articles that discuss geopolitical tensions, such as wars, military threats, and terrorist acts. What is relevant in our context, is that high GPR levels increase the likelihood of large-scale economic disruptions (Caldara and Iacoviello, 2022), and so can be interpreted as a measure of income risk faced by U.S. households.

We estimate the following regression model:

$$ser_t = \alpha + \beta_1 ln(y_t) + \beta_2 GPR_t + \beta_3 (ln(y_t))^2 + \sum_T \phi_T D_T + \epsilon_t,$$

where ser_t, denotes the share of the services in either the total nominal GDP or aggregate consumption; $ln(y_t)$ is the log of the yearly real gross domestic product per capita; GPR_t is the Geopolitical Risk index; $(ln(y_t))^2$ is the square of the log of the yearly real gross domestic product per capita; and D_T are twenty-year time fixed effects.

Unlike GDP volatility, which may be mechanically correlated with the structure of the economy through the size of the manufacturing sector (typically more volatile than the ser-

[13] Balleer, Duernecker, Forstner, and Goensch (forthcoming) document that households have biased subjective expectations when compared to actual probabilities of labor market transition. In our context, what is relevant is to capture the amount of risk that households actually perceive, as it determines precautionary savings and consumption choices.

Table 4: U.S. 1910-2008

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Value Added Share				Consumption Share			
$ln(y_t)$	0.117***	0.0560***	-1.187***	-1.271***	0.148***	0.0704***	-1.255***	-1.444***
	(0.00769)	(0.0201)	(0.206)	(0.306)	(0.00746)	(0.0233)	(0.208)	(0.329)
GPR_t	-0.000414***	-0.000378***	-0.000363***	-0.000360***	-0.000248***	-0.000210***	-0.000194***	-0.000190***
	(4.96e-05)	(4.73e-05)	(4.57e-05)	(4.51e-05)	(4.76e-05)	(4.80e-05)	(5.15e-05)	(4.78e-05)
$D_{1930-1949}$		0.0328**		0.0426***		0.0529***		0.0641***
		(0.0140)		(0.0132)		(0.0128)		(0.0117)
$D_{1950-1969}$		0.0209		0.0392**		0.0334*		0.0543***
		(0.0172)		(0.0177)		(0.0189)		(0.0197)
$D_{1970-1989}$		0.0626**		0.0546***		0.0915***		0.0824***
		(0.0247)		(0.0207)		(0.0283)		(0.0233)
$D_{1990-2008}$		0.123***		0.0696***		0.156***		0.0948***
		(0.0322)		(0.0223)		(0.0369)		(0.0262)
$ln(y_t)^2$			0.0695***	0.0722***			0.0748***	0.0825***
			(0.0107)	(0.0160)			(0.0109)	(0.0173)
Constant	-0.415***	0.106	5.670***	6.175***	-0.871***	-0.215	5.674***	6.714***
	(0.0773)	(0.175)	(0.990)	(1.451)	(0.0748)	(0.200)	(0.996)	(1.552)
Observations	99	99	99	99	99	99	99	99
R-squared	0.842	0.899	0.896	0.915	0.864	0.929	0.911	0.945
Time FE	NO	YES	NO	YES	NO	YES	NO	YES

Note: * p<0.1; ** p<0.05; *** p<0.01. Standard errors are in parentheses below coefficients.

vices sector), the GPR index is derived from textual analysis of news articles and reflects broader and external sources of uncertainty that are not directly related to sectoral composition. For this reason, and in contrast with the regressions in the previous section, we do not include the lagged manufacturing share among the control variables in these specifications.

Table 4 reports the regression results across different model specifications. Models (1) to (4) use the share of services in value added as the dependent variable, while models (5) to (8) use the share of services in consumption. Across all specifications, the coefficient on the GPR index is negative and statistically significant, indicating that higher levels of risk are associated with a lower share of services in both value added and consumption expenditure. This result supports the theoretical mechanism outlined earlier and align with the empirical results obtained with GDP volatility as a proxy for risk.

4.3 Cross-Sectional Evidence for the U.S.

In this section we investigate the relationship between income risk and the share of services in personal consumption expenditure using cross-sectional U.S. data. Specifically, we estimate household-level regressions where the dependent variable is the share of services in total consumption expenditure at a certain date, and the key explanatory variable is a measure of risk at the occupational level. The intuition is that occupations are heterogenous in income risk, so households in which the head is employed in a riskier occupation should display a lower share of services in consumption. To proxy occupational income risk, we employ alternative indicators including direct measures of labor market risk and exposure to routinization and automation. Control variables include age, educational attainment, household size, and other individual characteristics.

Let i denote the household and t the time period. The estimated regression is:

$$log(ser_{i,t}) = \alpha + \beta_1 \widehat{log(exp_{i,t})} + \beta_2\,(risk_{i,j,t}) + X_{i,t}\gamma + \mu_i + \varepsilon_{i,t} \tag{10}$$

where $log(ser_{i,t})$ is the logarithm of the service share total consumption expenditure; $\widehat{log(exp_{i,t})}$ is the logarithm of total consumption expenditure, which we instrument using the logarithm of income after taxes as in Boppart (2014), to control for endogeneity; $risk_{i,j,t}$ is a variable capturing income risk in occupation j, which might interact with carachteristics of household i; $X_{i,t}$ is a vector of control variables including household size, state of residence, education level and age of the household reference person (i.e. head of the household), μ_i are individual fixed effects and $\varepsilon_{i,t}$ idiosyncratic error term.

We estimate (10) using four different measures of occupational-level income risk. In

Model (1), $risk_{i,j,t}$ is defined as follows

$$risk^{(1)}_{i,j,t} = [log(exp_{i,t})] \frac{Weeks\ Unemp_{j,t-1}}{Weeks\ Worked_{j,t-1}},$$

where $log(exp_{i,t})$ is the logarithm of total expenditure of household i, $Weeks\ Unemp_{j,t-1}$ is the average of weeks unemployed in the previous year in occupation j and $Weeks\ Worked_{j,t-1}$ is the average of weeks worked in the previous year in occupation j.[14] The variable $risk^{(1)}_{i,j,t}$ measures the fraction of (log) consumption expenditure that the household would lose in case the head loses her current job in occupation j.

In Model (2) we use a similar variable to $risk^{(1)}_{i,j,t}$,

$$risk^{(2)}_{i,j,t} = [log(exp_{i,t})] \frac{Unemp\ Dur_{j,t}}{Weeks\ Worked_{j,t-1}},$$

but now we use the ratio of unemployment duration in the current year, $Unemp\ Dur_{j,t}$, to weeks worked in the previous year.[15]

In Model (3) we include the logarithm of the robot exposure measure at the occupational level developed by Cossu, Moro, and Rendall (2024).[16] Finally, in Model (4) we incorporate the logarithm of the occupational routine task index developed by Dorn (2009). Both indices are defined at the occupational level using the OCC1990dd and OCC1990 classification, which allows to aggregate them to the broader occupational categories reported in the CEX. Aggregation is performed using employment-weights, where weights reflect the relative occupational shares within each CEX group.[17]

Table 5 reports the results of instrumental variable regressions investigating the relationship between individual-level occupational income risk and the share of services in total consumption expenditure, using data from the Consumer Expenditure Survey (CEX) for the period 2014–2019. Following Boppart (2014) we instrument consumption expenditure with income level of the household. Our coefficient of interest is β_2, measuring the effect of occupational income risk on the share of services. In line with our theoretical model, all four specifications display a negative and statistically significant relationship between income

[14]We use IPUMS-CPS variables WKSWORK1, number of weeks worked in the previous year and WKSUNEM1, number of weeks unemployed in the previous year.

[15]We use IPUMS-CPS variables WKSWORK1, number of weeks worked in the previous year and DURUNEMP: duration of the current unemployment spell (in weeks).

[16]This measure combines the O*NET taxonomy of Intermediate Work Activities (IWAs) with a classification of robots applications provided by the International Federation of Robotics (IFR). Exposure to robotization is then defined as the share of IWAs within each occupation that are performable by existing robots.

[17]Appendix A provides details on the data and the data treatments used.

risk at the occupational level and the service share of the household. This indicates that individuals employed in occupations characterized by greater income risk allocate a smaller proportion of their consumption to services, likely reflecting a precautionary savings behavior. These findings are consistent with the time series evidence reported in the previous sections.

Table 5: Cross-sectional evidence for the U.S.

	(1)	(2)	(3)	(4)
$\widehat{log(exp_{i,t})}$	0.156^{***}	0.152^{***}	0.123^{***}	0.123^{***}
	(0.005)	(0.005)	(0.005)	(0.005)
$risk_{i,t}$	-0.114^{***}	-0.119^{***}	-0.005^{**}	-0.004^{***}
	(0.003)	(0.003)	(0.002)	(0.002)
$Residence\ indicators$	YES	YES	YES	YES
$Family\ size\ indicators$	YES	YES	YES	YES
$Ref.person\ controls$	YES	YES	YES	YES
$Data$	CEX	CEX	CEX	CEX
$Sample\ years$	$2014-2019$	$2014-2019$	$2014-2019$	$2014-2019$
$Method$	IV	IV	IV	IV
$Observations$	$56,103$	$56,103$	$56,103$	$56,103$

Note: * p<0.1; ** p<0.05; *** p<0.01. Standard errors are in parentheses below coefficients. All regressions include quarter fixed effects.

5 Cross-Country

Having established the empirical relevance of our mechanism in the context of the U.S., we now explore whether a similar relationship between GDP volatility and sectoral composition holds in a cross-country setting. More specifially, we investigate the potential role of our mechanism in shaping the phenomenon of *premature de-industrialization,* as defined in Rodrik (2016). In a nutshell, this phenomenon shows that at a given level of GDP per-capita, countries undergoing structural transformation later in time display a smaller share of manufacturing value added compared to countries that underwent structural transformation earlier in time. In other words, the share of manufacturing in countries experiencing premature de-industrialization grows less and starts declining at lower levels of income with respect to countries that industrialized (and de-industrialized) earlier.

Our theory proposes a potential channel to explain the phenomenon, which complements

Table 6: Cross-Country estimates 1950-2010

	South America		Asia		Asia Exporters		Asia Non Exporters		All Countries	
	(2)	(6)	(2)	(6)	(2)	(6)	(2)	(6)	(2)	(6)
$ln(y_t)$	-0.0945***	0.936***	0.0437***	0.138**	0.0556***	0.365***	0.0939***	-0.357*	0.0325***	0.130**
	(0.0177)	(0.245)	(0.00567)	(0.0610)	(0.0112)	(0.0594)	(0.00876)	(0.207)	(0.00650)	(0.0543)
$man_{i,t-5}$	-0.112*	0.613	-0.691***	-1.271***	-0.952***	-1.605***	-0.290***	0.116	-0.354***	0.238
	(0.0621)	(0.422)	(0.0442)	(0.264)	(0.108)	(0.314)	(0.0483)	(0.316)	(0.0461)	(0.233)
$\sigma_{i,t:-15}$	-1.047***	-1.041***	-0.273	-0.0549	-0.0511	0.816***	-3.041***	-2.102***	-0.765***	-0.652***
	(0.258)	(0.263)	(0.184)	(0.248)	(0.218)	(0.269)	(0.483)	(0.534)	(0.198)	(0.217)
$D_{1970-1979}$	0.0182**	0.0119*	0.00933	0.00773	0.0116	-0.00450	-0.0118	-0.00212	0.00120	-0.00352
	(0.00847)	(0.00721)	(0.00729)	(0.00863)	(0.00778)	(0.00978)	(0.00995)	(0.0103)	(0.00689)	(0.00634)
$D_{1980-1989}$	0.0258***	0.0223***	0.0700***	0.0692***	0.0814***	0.0570***	0.00543	0.0188	0.0236***	0.0181***
	(0.00931)	(0.00851)	(0.00880)	(0.0109)	(0.00946)	(0.0130)	(0.0150)	(0.0162)	(0.00743)	(0.00689)
$D_{1990-1999}$	0.0991***	0.0899***	0.112***	0.110***	0.121***	0.0932***	0.0151	0.0393*	0.0712***	0.0634***
	(0.0118)	(0.0120)	(0.0104)	(0.0125)	(0.0104)	(0.0136)	(0.0191)	(0.0211)	(0.00850)	(0.00781)
$D_{2000-2010}$	0.118***	0.110***	0.136***	0.130***	0.149***	0.117***	0.0338	0.0490**	0.0827***	0.0763***
	(0.0137)	(0.0140)	(0.0117)	(0.0141)	(0.0124)	(0.0151)	(0.0217)	(0.0222)	(0.00985)	(0.00906)
$ln(y_t)^2$		-0.0557***		-0.00494		-0.0170***		0.0240**		-0.00540*
		(0.0136)		(0.00314)		(0.00323)		(0.0109)		(0.00302)
$man^2_{i,t-5}$		-1.048*		0.726**		0.937**		-0.393		-0.835***
		(0.556)		(0.327)		(0.409)		(0.387)		(0.301)
Constant	1.400***	-3.467***	0.232***	-0.103	0.206***	-1.075***	-0.239***	1.733*	0.286***	-0.250
	(0.155)	(1.110)	(0.0411)	(0.277)	(0.0623)	(0.262)	(0.0717)	(0.902)	(0.0477)	(0.241)
Observations	291	291	331	331	214	214	117	117	622	622
R-squared	0.738	0.756	0.911	0.913	0.909	0.917	0.963	0.965	0.862	0.865
Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Country FE	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: * p<0.1; ** p<0.05; *** p<0.01. Standard errors are in parentheses below coefficients. Numbers (2) and (6) refer to the corresponding econometric model in Table 1.

those in Huneeus and Rogerson (2024) and Sposi, Yi, and Zhang (2024). Countries that developed earlier like the U.S. underwent the process of structural transformation at higher levels of income volatility with respect to countries that went throught the same process later in time, as suggested by Figure 1. These differences in GDP volatility might be due to more developed international capital markets and trade integration in the second half of the century, something that the U.S. did not benefit from during the rise of its manufacturing sector, or to major events like the two World Wars. Lower income risk at similar income levels reduces precautionary savings and accelerates the shift away from manufacturing as predicted by the theory in section 3. Rodrik (2016) documents that Latin American and Asian countries were affected by premature de-industrialization. Following these findings, we then focus on Latin America and Asia and study whether risk played a role in shaping observed structural transformation.

To test our hypothesis, we use data from the GGDC 10-Sector Database and Penn World Table version 10.01, both provided by Groningen University.[18] We use data from 1950 to 2010 for countries in South America and Asia. We estimate a regression similar to that used for the U.S. time-series analysis, in which risk is proxied by past GDP volatility, but now we exploit the panel dimension of the dataset and include country fixed effects. In addition, we focus on our preferred specifications from Table 2, (2) and (6), which include time fixed effects.

$$ser_{i,t} = \alpha_i + \beta_1 ln(y_{i,t}) + \beta_2 man_{i,t-5} + \beta_3 \sigma_{i,t:-15} + \beta_4 (ln(y_{i,t}))^2 + \beta_5 (man_{i,t-5})^2 + \sum_T \phi_T D_T + \sum_T \gamma_T F_T + \epsilon_{i,t}.$$

Table 6 reports the results. The first two columns report South America. Both specifications find a statistically significant negative coefficient of income volatility on the share of services. This suggests that the smaller level of risk faced by South American countries with respect to that measured for the U.S. might have determined part of the premature de-industrialization observed.

The third and the fourth columns report results for Asia. In this case the coefficient on volatility is still negative in both models, but not statistically significant. This suggests that for these countries volatility is a weaker driver of structural transformation, with the other regressors playing a more important role. To assess whether openness to trade plays a central role in the results for Asia, as suggested in Rodrik (2016), we split the sample of Asian countries into two groups: manufacturing exporters and non-exporters. We do not perform this split for South America, as all countries in that region are classified as non-exporters, making the sample composition effectively unchanged. Following Rodrik (2016),

[18] See Appendix A for details.

we classify countries as manufacturing exporters if the share of manufacturing in total exports exceeds its share in imports, using bilateral trade data from the World Bank.[19] We then re-estimate our baseline regressions separately for the two subsamples of Asian countries.[20] The results, reported in columns 5-8 in Table 6, offer strong support for our mechanism and are consistent with Rodrik's broader argument. Among Asian manufacturing exporters, we find either no significant relationship between GDP volatility and the share of services (model 2) or even a statistically significant and positive one (model 6), in line with the idea that external demand and trade integration buffer the domestic economy against income risk, thus canceling the role of precautionary savings in shaping structural change. In contrast, among non-exporting Asian economies, the relationship between GDP volatility and the service sector share is negative and statistically significant, closely mirroring the pattern observed in South America, and even with a larger absolute value of the coefficient. These findings suggest that openness modulates the strength of the volatility–structural change link: when economies are less integrated into global manufacturing trade, GDP volatility exerts a stronger dampening effect on the expansion of services, via higher precautionary savings and reduced domestic demand.

Finally, columns 9 and 10 pool all South-American and Asian countries together. In this case the relationship is strongly significant and negative, suggesting that when considering the whole group of premature de-indistrializers, risk is an important factor determining the pattern of structural transformation.

6 Premature de-industrializers or Late industrializer?

In this section, we present evidence on the differences in income volatility between premature de-industrializers and the United States. Table 7 reports GDP volatility - measured as percentage deviations from an HP-filtered trend - for the countries in our sample for which we have the entire post-WWII period (1953-2010) and for the historical data of the U.S. Both South American and Asian economies in the post–World War II period exhibit GDP volatility levels that are more comparable to those of the U.S. during the same period than to the U.S. levels observed in the pre-war era. This finding is consistent with the visual

[19] More specifically, from the World Bank Development Indicators we use data on Manufactures exports (% of merchandise exports) and Manufactures imports (% of merchandise imports). We compute, for each country in our sample, the average export and import shares over the period 1994–2010, starting in 1994 due to incomplete data in earlier years. Based on this criterion, we identify China, India, Korea, Philippines, and Thailand as manufacturing exporters.

[20] The sample for South America is basically the same if we split countries by export: all countries are non exporters except Mexico, and the sample is too short to run a single country regression. So the split is meaningful only for Asia.

Table 7: Share of Services and Volatility

Country	Years	Initial Share	Final Share	Mean Volatility
U.S. (Mad)	1910-1952	0.49	0.58	0.049
U.S. No WWII (Mad)	1910-1952	0.49	0.58	0.043
U.S. (Mad)	1953-2008	0.58	0.79	0.023
U.S.	1953-2010	0.52	0.77	0.014
ARG	1953-2010	0.50	0.44	0.032
CHL	1953-2010	0.64	0.55	0.032
COL	1953-2010	0.42	0.53	0.014
CRI	1953-2010	0.43	0.59	0.020
MEX	1953-2010	0.48	0.53	0.022
South America Average		0.49	0.53	0.024
THA	1953-2010	0.33	0.34	0.030
TWN	1953-2010	0.36	0.60	0.017
CHN	1953-2010	0.25	0.34	0.034
KOR	1953-2010	0.40	0.52	0.021
IND	1953-2010	0.24	0.45	0.020
Asia Average		0.32	0.45	0.024

Note: First three lines are computed using Maddison data from Herrendorf, Rogerson, and Valentinyi (2014). All the other measures are from the GGDC 10-Sector Database and Penn World Table version 10.01. U.S. No WWII (Mad) excludes observations in the ten years around WWII, from 1941 to 1950. We split the sample in 1952/53 because that is the peak of the manufacturing share in the U.S.

evidence presented in Figure 1.

These observations suggest that an interpretation of our theory is that the U.S. can be viewed as a *late* (in income*) industrializer* relative to the South American and Asian economies, due to the fact that it experienced particularly high volatility during the period of expansion in the manufacturing sector, which delayed its transition toward a service-based economy. The estimates in Table 2 allow to perform a counterfactual exercise. Suppose that U.S. GDP volatility in the pre-WWII period had been equal to its post-WWII level. This would imply a reduction in volatility of 0.026 (i.e., 0.023 - 0.049). Given the estimated coefficient on volatility in model (6) of Table 2 is -0.895, this lower volatility would translate into an average increase of 0.0233 (= -0.026 $\times$ -0.895) in the U.S. share of services during the pre-WWII period. To put this into perspective, Figure 1 shows that the average deviation of the historical U.S. share of services (yellow triangles) from the estimated trend (depicted by the black line) is 0.075. Hence, our mechanism accounts for about 31% of the observed difference in the share of services between the U.S. and the group of premature de-industrializers.

7 Conclusions

This paper presents a mechanism linking micro-level risk to macroeconomic structural composition, thus providing a novel explanation for how income volatility influences structural change. In a two-sector model with non-homothetic preferences, we show that greater income risk leads to increased precautionary savings, which reduce consumption expenditure and lower the share of services in both consumption and value added at a given GDP level.

We provide an empirical analysis supporting the mechanism, by exploiting using U.S. historical time-series and different measures of risk, and U.S. cross-sectional consumption data interacted with income risk at the occupational level. The analysis for the U.S. provides robust support for the link between risk and the structural composition of the U.S. economy. This suggests that, had the U.S. experienced lower income risk along the development path, it would also have experienced a faster transition towards the services sector.

Our results offer a novel explanation for premature de-industrialization, which is based on income risk. Countries that industrialized later in time, such as those in South America and parts of Asia, faced lower income risk with respect to early industrializers like the U.S. In this view, the U.S. can be considered a late industrializer with respect to those countries. Lower risk in Latian American and Asian countries reduces precautionary savings, raises consumption, and shifts demand toward services due to non-homothetic preferences. This leads to an increase in the services share at lower income levels with respect to the U.S. Among Asian countries, the empirical evidence shows a strong negative link between GDP volatility and service shares in non-exporting countries, but not in exporting ones, confirming that our mechanism is especially relevant in a closed economy setting.

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Appendix

A Data

A.1 U.S. Historical Data from 1900 to 2008

We rely on the historical time series compiled by Herrendord et al. (2014), which integrate data from Carter et al. (2006) for the pre-1930 period with data from the Bureau of Economic Analysis (BEA) starting in 1929. This dataset is further supplemented with GDP per capita estimates from Maddison (2010), covering the years 1800–2000. The dataset provides historical U.S. data on sectoral value-added (for agriculture, manufacturing, and services) and personal consumption expenditures, both expressed in millions of U.S. dollars at current prices. Although the GDP per capita series begins in 1800, sectoral value-added data - necessary for computing sectoral shares - are available on a continuous yearly basis only from 1909 onward, while personal consumption expenditure data are available starting in 1900. Given that our volatility indicator requires a 15-year backward window at each year t, we use Maddison's real GDP per capita series (expressed in 1990 international Geary-Khamis dollars) beginning in 1894. Furthermore, we compute the sectoral shares - Agriculture, Manufacturing, and Services - on both the value-added and consumption sides, defined as the ratio of each sector's value to the total of the three sectors. The value-added shares cover the period from 1909 to 2008, while the consumption shares span the years 1900 to 2008.

A.2 GGDC 10-Sector Database

The GGDC 10-Sector Database, provided by the Groningen Growth and Development Centre (GGDC), provides annual data for 42 countries across Africa, Asia, Europe, Latin America, the Middle East, and North America. Covering the period from 1947 to 2013, the database includes information on persons employed, gross value added at current national prices, and gross value added at constant 2005 national prices for 10 broad economic sectors. For the purpose of our analysis, we aggregate the 10 sector into three broad categories using data from the period 1950 to 2010. Agriculture corresponds to 'Agriculture, hunting, forestry, and fishing' (AtB). Manufacturing includes 'Mining and quarrying' (C), 'Manufacturing' (D), 'Electricity, gas, and water supply' (E), and 'Construction' (F). Services encompass 'Wholesale and retail trade, hotels, and restaurants' (GtH), 'Transport, storage, and communication' (I), 'Finance, insurance, real estate, and business services' (JtK), and 'Community, social, and personal services' (OtP) and 'Government services' (LtN).

We compute sectoral shares - Agriculture, Manufacturing, and Services - as the ratio of each aggregated sector's nominal value added to the total of the three aggregated sectors. In the cross-country regressions we use the following economies. For ASIA: China (CHN, 1952 to 2010), Indonesia (IDN, 1960 to 2010), India (IND, 1950 to 2010), Korea Republic (KOR, 1953 to 2010), Malaysia (MYS, 1970 to 2010), Philippines (PHL, 1971 to 2010), Thailand (THA, 1951 to 2010), Taiwan (1951 to 2010). For SOUTH AMERICA: Argentina (ARG, 1950 to 2010), Bolivia (BOL, 1958 to 2010), Brazil (BRA, 1990 to 2010), Chile (CHL, 1950 to 2010), Colombia (COL, 1950 to 2010), Costa Rica (CRI, 1950 to 2010), Mexico (MEX, 1950 to 2010). We exclude Perù (PER) and Venezuela (VEN) from South-American countries as they either display negative value added for some sector (Perù) or display a structural change that is not comparable to the rest of coutries (Venezuela). We acknowledge, however, that most results for South-America hold with the inclusion of Venezuela.

A.3 GGDC PWT 10.01

The Penn World Table (PWT) version 10.01, provided by the Groningen Growth and Development Centre (GGDC), is a comprehensive database covering 183 countries from 1950 to 2019. It provides detailed information on relative income levels, output, input, and productivity, including variables such as real GDP (both expenditure-side and output-side) at chained and constant purchasing power parity (PPP), employment levels, human capital indices, capital stock, total factor productivity (TFP), and exchange rates. For our analysis, we constructed a GDP per capita series by dividing the "Real GDP at constant 2017 national prices - RGDPNA" (in millions of 2017 US dollars) by "Population" (in millions).

A.4 Volatility Indicator based on GDP and Aggregate Consumption

To construct our volatility indicator, we apply the Hodrick–Prescott (HP) filter to the logarithmic transformation of the real GDP per capita series in order to isolate its cyclical component. The smoothing parameter is set to $\lambda = 6.25$, following Ravn and Uhlig (2002) for annual data, which adjusts the filter to preserve its frequency response characteristics. To capture the volatility of real GDP and real consumption per capita in each year t, we compute a rolling standard deviation of the log-deviations from trend over the previous 15 years - from $t-16$ to $t-1$. This is our measure of volatility $\sigma_{t:-15}$ at time t that we use in Tables 2 and 6 for GDP and Table 3 for consumption.

A.5 Geopolitical Risk index

As an alternative proxy for macroeconomic uncertainty, we use the Historical Geopolitical Risk (GPR) Index developed by Caldara and Iacoviello (2022). This index quantifies the intensity of geopolitical tensions using a news-based methodology that measures the frequency of articles discussing wars, terrorism, and international crises in major newspapers (e.g., New York Times, Washington Post, Chicago Tribune). Their historical index is constructed from 1900 onward at monthly frequency, based on the share of articles containing pre-specified keywords related to geopolitical threats and acts. To integrate this indicator into our annual U.S. database, we compute, for each year, the simple average of the twelve monthly historical GPR values, thus generating a yearly series of geopolitical risk consistent with the temporal resolution of our main dataset.

A.6 Cross-sectional data for the U.S.

Our empirical analysis relies on data from the U.S. Consumer Expenditure Survey (CEX), which provides detailed information on household consumption, income, and demographic characteristics. We use the quarterly consumer unit files covering the period 2014–2019, focusing on cross-sectional variation in individual spending behavior.

The dependent variable, the share of services in total consumption, is constructed following the methodology proposed by Boppart (2014). In particular, we define service consumption to include the following categories: food away from home; shelter; utilities, fuels, and public services; other vehicle expenses; public transportation; health care; personal care; education; cash contributions; personal insurance; and pensions. The dataset also includes a wide range of individual-level control variables, such as age, education, household size, and geographical identifiers. These variables allow us to control for key socio-demographic factors that may influence consumption patterns independently of income risk.

To measure occupational-level income risk, we link CEX respondents to a set of external indicators based on their reported occupation. These indicators capture different dimensions of labor market uncertainty and include both technology-related proxies, such as direct measures of employment instability - as the duration of unemployment spells and the number of weeks unemployed in the previous year - and exposure to automation and routinization.

Among the indicators used to proxy occupational income risk, $risk^{(1)}_{i,j,t}$ and $risk^{(2)}_{i,j,t}$, both are computed by OCC1990 occupation code, and subsequently aggregated to the CEX occupational classification using ASEC weights (asecwt). For $risk^{(3)}_{i,j,t}$ we use a robot exposure measure developed by Cossu, Moro, and Rendall (2024). For $risk^{(4)}_{i,j,t}$, we use the routine task intensity index developed by Dorn (2009). Both indices are defined at the occupational level using the OCC1990dd and OCC1990 classification, which allows us to aggregate them to the broader occupational categories reported in the CEX. Aggregation is performed using employment-weights, where weights reflect the relative occupational shares within each CEX group.

To align the occupational-level risk measures with the CEX classification, we map all indicators with those derived from CPS data and those based on the routine and robotization indices to the occupational categories reported in the CEX. This is done using a consistent crosswalk between OCC1990 and OCC1990dd codes. We then group occupations into the 14 major occupational categories defined by the CEX, excluding the 15th category, which refers to members of the armed forces.[21]

B Model Extensions

In this Appendix we extend the model in two dimensions. First, we consider that the investment good is also a composite of manufacturing and services. Second, we consider Epstein-Zin-Weil preferences, that allow to discipline the intertemporal elasticity of substitution and risk aversion separately.

B.1 Model with composite investment

B.1.1 Static Consumtion Problem

First, the households maximize the consumption index, as in the basic model, at each t

$$\max_{c_{m,t}, c_{s,t}} c_t = [\omega_m^{1/\varepsilon} c_{m,t}^{\frac{\varepsilon-1}{\varepsilon}} + \omega_s^{1/\varepsilon} (c_{s,t} + s)^{\frac{\varepsilon-1}{\varepsilon}}]^{\frac{\varepsilon}{\varepsilon-1}},$$

[21] We define: Manager, Professional administrators occ1990dd>=3 & occ1990dd<=37; Teacher occ1990dd>=154 & occ1990dd<=159; Professional occ1990dd>=43 & occ1990dd<=153 & occ1990dd>=160 & occ1990dd<=200 ; Administrative support occ1990dd>=303 & occ1990dd<=389; Sales retail occ1990dd>=274 & occ1990dd<=283; Sales, Business Goods and Services: occ1990dd>=243 & occ1990dd<=258; occ1990dd>=433 & occ1990dd<=444; occ1990dd>=445 & occ1990dd<=447; Technician Service: occ1990dd>=203 & occ1990dd<=235; Protective Service: occ1990dd>=415 & occ1990dd<=427; Private Household Service: occ1990dd>=405 & occ1990dd<=408; occ1990dd>=448 & occ1990dd<=455; Other Service: occ1990dd>=457 & occ1990dd<=458; occ1990dd>=459 & occ1990dd<=467; occ1990dd>=469 & occ1990dd<=472; occ1990dd==468; Machine or Transportation Operator, Laborer: occ1990dd>=503 & occ1990dd<=549; occ1990dd>=703 & occ1990dd<=799; occ1990dd>=803 & occ1990dd<=889; Construction Workers, Mechanics: occ1990dd>=558 & occ1990dd<=599; occ1990dd>=628 & occ1990dd<=699; Farming: occ1990dd>=473 & occ1990dd<=489; Forestry, Fishing, Groundskeeping: occ1990dd>=496 & occ1990dd<=498.

subject to an expenditure constraint

$$p_{m,t}c_{m,t} + p_{s,t}c_{s,t} = \bar{w},$$

This problem delivers, as showed in the text, the following solution

$$c_{m,t} = \left(\frac{p_{m,t}}{p_{s,t}}\right)^{-\varepsilon} \frac{\omega_m}{\omega_s}(c_{s,t} + s),$$

$$c_{s,t} = \frac{\frac{\bar{w}}{p_{m,t}}\left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon} - \frac{\omega_m}{\omega_s}s}{\frac{\omega_m}{\omega_s} + \left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon-1}},$$

where $\bar{w}$ is an exogenous expenditure level.

Also, the first order conditions allow to show that

$$(c_t)^{\frac{1}{\varepsilon}} \omega_g^{1/\varepsilon}(c_{g,t})^{-\frac{1}{\varepsilon}} = \lambda p_{g,t}$$

$$(c_t)^{\frac{1}{\varepsilon}} \omega_s^{1/\varepsilon}(c_{s,t} + s)^{-\frac{1}{\varepsilon}} = \lambda p_{s,t}$$

$$p_{m,t}c_{m,t} + p_{s,t}c_{s,t} = p_t c_t - p_{s,t}s.$$

where

$$p_t = \left[\omega_g \left(p_{m,t}\right)^{1-\varepsilon} + \omega_s \left(p_{s,t}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}$$

is the price index of the consumption index c_t.

B.1.2 Static Investment Problem

The households maximize the following static investment index at each $t = 0$

$$\underset{i_{m,t},i_{s,t}}{max}\, k_1 = \left[\omega_{i,m}^{1/\varepsilon}\left(i_{m,t}\right)^{\frac{\varepsilon-1}{\varepsilon}} + \omega_{i,s}^{1/\varepsilon}\left(i_{s,t}\right)^{\frac{\varepsilon-1}{\varepsilon}}\right]^{\frac{\varepsilon}{\varepsilon-1}},$$

subject to an expenditure constraint

$$p_{m,t}i_{m,t} + p_{s,t}i_{s,t} = \bar{w}_i,$$

This problem delivers as solution

$$i_{m,t} = \left(\frac{p_{m,t}}{p_{s,t}}\right)^{-\varepsilon} \frac{\omega_{i,g}}{\omega_{i,s}}(i_{s,t}).$$

and

$$i_{s,t} = \frac{\frac{\bar{w_i}}{p_{m,t}}\left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon}}{\frac{\omega_{i,g}}{\omega_{i,s}} + \left(\frac{p_{m,t}}{p_{s,t}}\right)^{\varepsilon-1}},$$

Also, the first order conditions allow to show that

$$p_{g,t}i_{m,t} + p_{s,t}i_{s,t} = p_{k,t}k_{t+1},$$

where

$$p_{k,t} = \left[\omega_{i,g}\left(p_{g,t}\right)^{1-\varepsilon} + \omega_{i,s}\left(p_{s,t}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}. \tag{11}$$

is the price index of the investment index purchased at time t, k_{t+1}.

B.1.3 Dynamic problem with CRRA preferences

By using the above results from the static problem, we can then rewrite the original problem as

$$\max_{\{c_0,c_1(q),k_1\}}\left[\frac{c_0^{1-\sigma}}{1-\sigma} + \beta E\left[\frac{c_1(q)^{1-\sigma}}{1-\sigma}\right]\right],$$

subject to

$$p_0c_0 + p_{k,0}k_1 = p_{m,0}A_{m,0}k_0 + p_{s,0}s,$$

$$p_1(q)c_1(q) = p_{m,1}(q)A_{m,1}(q)k_1 + p_{s,1}(q)s,$$

where we used that $p_{m,0}A_{m,0}k_0 = r_0k_0$ and $p_{m,1}(q)A_{m,1}(q)k_1 = r_1(q)k_1$ for each state of the world.

$$max\left[\frac{c_0^{1-\sigma}}{1-\sigma} + \beta E\left[\frac{c_1^{1-\sigma}}{1-\sigma}\right]\right],$$

subject to

$$c_0 = \frac{p_{g,0}a_{g,0}k_0 + p_{s,0}s - p_{k,0}k_1}{p_0} = \frac{a_{g,0}k_0 + \frac{p_{s,0}}{p_{g,0}}s - \frac{p_{k,0}}{p_{g,0}}k_1}{p_0/p_{g,0}},$$

$$c_1(q) = \frac{p_{g,1}(q)a_1(q)k_1 + p_{s,1}(q)s}{p_1(q)} = \frac{a_1(q)k_1 + \frac{p_{s,1}(q)}{p_{g,1}(q)}s}{p_1(q)/p_{g,1}(q)}.$$

We can substitute for c_0 and $c_1(q)$ in the problem to get

$$\underset{k_1}{max}\left[\frac{\left(\frac{a_{g,0}k_0+\frac{p_{s,0}}{p_{g,0}}s-\frac{p_{k,0}}{p_{g,0}}k_1}{p_0/p_{g,0}}\right)^{1-\sigma}}{1-\sigma}+\beta E\left[\frac{\left(\frac{a_1(q)k_1+\frac{p_{s,1}(q)}{p_{g,1}(q)}s}{p_1(q)/p_{g,1}(q)}\right)^{1-\sigma}}{1-\sigma}\right]\right],$$

The first order condition with respect to $k_1$is now

$$u'\left(\frac{a_{g,0}k_0+\frac{p_{s,0}}{p_{g,0}}s-\frac{p_{k,0}}{p_{g,0}}k_1}{p_0/p_{g,0}}\right)\frac{p_{k,0}}{p_0}=\beta E\left[u'\left(\frac{p_{g,1}(q)a_1(q)k_1+p_{s,1}(q)s}{p_1(q)}\right)\frac{p_{g,1}(q)a_1(q)}{p_1(q)}\right].$$

Considering the derivative of the CRRA $U'=c^{-\sigma}$

$$\left(\frac{a_{g,0}k_0+\frac{p_{s,0}}{p_{g,0}}s-\frac{p_{k,0}}{p_{g,0}}k_1}{p_0/p_{g,0}}\right)^{-\sigma}\frac{p_{k,0}}{p_0}=\beta E\left[\left(\frac{p_{g,1}(q)a_1(q)k_1+p_{s,1}(q)s}{p_1(q)}\right)^{-\sigma}\frac{p_{g,1}(q)a_1(q)}{p_1(q)}\right].$$

The Euler equation for the problem then becomes

$$\left(\frac{a_{g,0}k_0+\frac{p_{s,0}}{p_{g,0}}s-\frac{p_{k,0}}{p_{g,0}}k_1}{p_0/p_{g,0}}\right)^{-\sigma}\frac{1}{p_0/p_{k,0}}=\beta E\left[\left(\frac{a_1(q)k_1+\frac{p_{s,1}(q)}{p_{g,1}(q)}s}{p_1(q)/p_{g,1}(q)}\right)^{-\sigma}\frac{a_1(q)}{p_1(q)/p_{g,1}(q)}\right],\quad (12)$$

where

$$p_{s0}/p_{g0}=a_{g0}/a_{s0},$$

$$p_{s1}(q)/p_{g1}(q)=a_{g1}(q)/a_{s1}(q),$$

$$p_0/p_{m,0}=\left[\omega_{c,m}+\omega_{c,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}},$$

$$p_{k,0}/p_{m,0}=\left[\omega_{i,m}+\omega_{i,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}},$$

$$p_0/p_{k,0}=\frac{\left[\omega_{c,m}+\omega_{c,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}}{\left[\omega_{i,m}+\omega_{i,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}},$$

$$p_1(q)/p_{g1}(q)=[\omega_g 1+\omega_s\left[\frac{p_{s1}(q)}{p_{g1}(q)}\right]^{1-\varepsilon}]^{\frac{1}{1-\varepsilon}}.$$

In Figure 4 we report investment, consumption expenditure and structural composition in both consumption and value added in the model considering the parametrization reported in Table 8. We can observe the same qualitative effects of volatility all variables as in the benchmark model in Figure 2.[22]

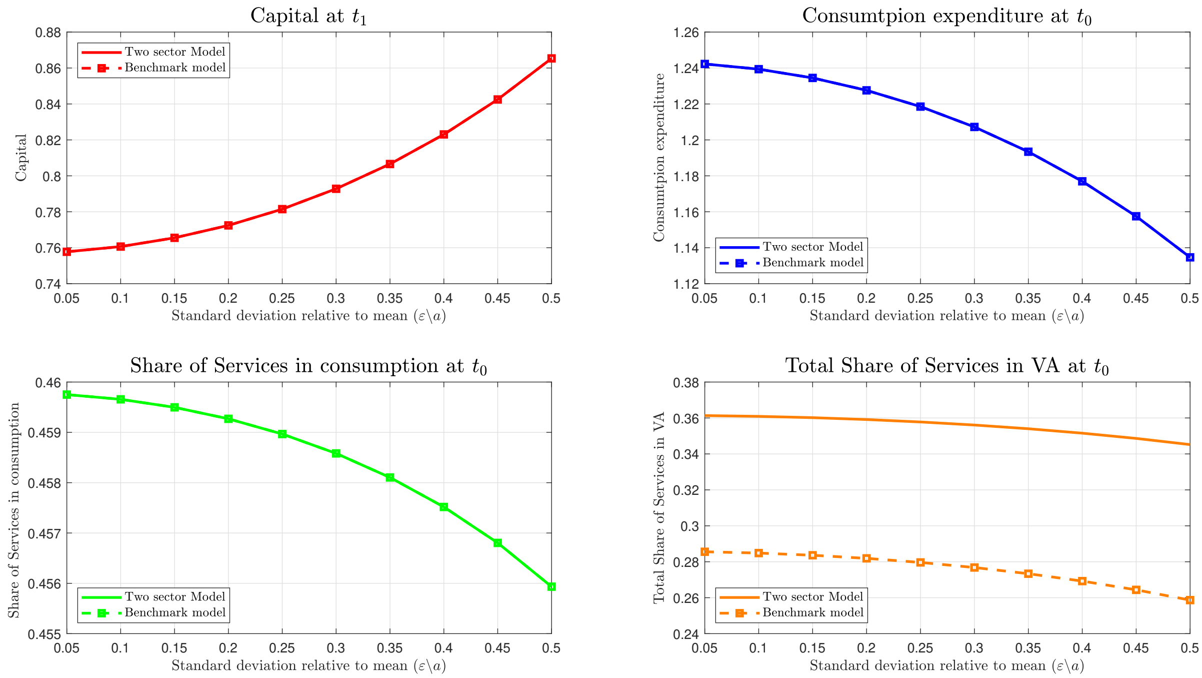

Figure 4: Comparison between Benchmark and Two sector model

B.2 Dynamic problem with Epstein-Zin-Weil preferences

Using Epstein-Zin-Weil preferences lifetime utility becomes

$$V_0 = \left\{ (1-\beta)\, c_0^{1-\frac{1}{\psi}} + \beta E\left[c_1(q)^{1-\gamma}\right]^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right\}^{\frac{1}{1-\frac{1}{\psi}}},$$

subject to

$$c_0 = \frac{a_{m,0}k_0 + \frac{p_{s,0}}{p_{m,0}}s - \frac{p_{k,0}}{p_{m,0}}k_1}{p_0/p_{m,0}},$$

[22] To understand why also the quantitative pattern is the same as in the benchmark model, note that the only difference in the dynamic problems is the price of investment in period 0. This is the price of goods in the benchmark model in the text while it is a composite of goods and services, given by (11) in the model with composite investment. However, given the assumption on initial TFP levels, the two prices in the two models have the same value, making the solution of the dynamic problem the same in the two cases.

Table 8: Models parameters

Parameters	**Symbol**	**Value**
Risk aversion	σ	3.5
Discount factor	β	0.95
Minimum required consumption in services	s	0.1
Elasticity of substitution between manufacturing and services	ε	0.5
Initial capital	k_0	1
Stone-Geary consumption weights	$\omega_{c,m} = \omega_{c,s}$	0.5
Stone-Geary investment weight – manufacturing	$\omega_{i,m}$	0.8
Stone-Geary investment weight – services	$\omega_{i,s}$	0.2
Total factor productivities (at $t = 0$)	$a_{m,0} = a_{s,0}$	2
Probabilities of each states	$P_1 = P_2 = P_3 = P_4$	0.25

$$c_1(q) = \frac{a_{m,1}(q)k_1 + \frac{p_{s,1}(q)}{p_{m,1}(q)}s}{p_1(q)/p_{m,1}(q)},$$

where β is the discount rate, γ is the coefficient of relative risk aversion and ψ is the elasticity of inter-temporal substitution.

Thus we can rewrite the original problem as

$$\underset{c_0,c_1,}{max} V_0 = \underset{c_0,c_1,}{max} \left\{ (1-\beta)\, c_0^{1-\frac{1}{\psi}} + \beta E\left[c_1(q)^{1-\gamma}\right]^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right\}^{\frac{1}{1-\frac{1}{\psi}}}$$

subject to

$$c_0 = \frac{a_{m,0}k_0 + \frac{p_{s,0}}{p_{m,0}}s - \frac{p_{k,0}}{p_{m,0}}k_1}{p_0/p_{m,0}},$$

$$c_1(q) = \frac{a_{m,1}(q)k_1 + \frac{p_{s,1}(q)}{p_{m,1}(q)}s}{p_1(q)/p_{m,1}(q)},$$

and substituting into the V_0 we obtain

$$\max_{k_1} \left\{ (1-\beta)\left(\frac{a_{m,0}k_0 + \frac{p_{s,0}}{p_{m,0}}s - \frac{p_{k,0}}{p_{m,0}}k_1}{p_0/p_{m,0}} \right)^{1-\frac{1}{\psi}} + \beta E\left[\left(\frac{a_{m,1}(q)k_1 + \frac{p_{s,1}(q)}{p_{m,1}(q)}s}{p_1(q)/p_{m,1}(q)} \right)^{1-\gamma} \right]^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right\}^{\frac{1}{1-\frac{1}{\psi}}}$$

The first order condition with respect to k_1 gives

$$\begin{gathered} 0 = \frac{1}{1-\frac{1}{\psi}} \left\{ (1-\beta)\, c_0^{1-\frac{1}{\psi}} + \beta E\left[c_1(q)^{1-\gamma}\right]^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right\}^{\frac{1}{1-\frac{1}{\psi}}-1} \cdot \\ \cdot \left\{ -\frac{1}{p_0/p_{k,0}}(1-\beta)\left(1-\frac{1}{\psi}\right) c_0^{-\frac{1}{\psi}} + \beta\left(1-\frac{1}{\psi}\right) \cdot E\left[c_1(q)^{1-\gamma}\right]^{\frac{1-\frac{1}{\psi}}{1-\gamma}-1} \cdot E\left[\frac{a_{m,1}(q)}{p_1(q)/p_{m,1}(q)} c_1(q)^{-\gamma} \right] \right\}, \end{gathered} \tag{13}$$

where

$$c_0 = \frac{a_{m,0}k_0 + \frac{p_{s,0}}{p_{m,0}}s - \frac{p_{k,0}}{p_{m,0}}k_1}{p_0/p_{m,0}},$$

$$c_1(q) = \frac{a_{m,1}(q)k_1 + \frac{p_{s,1}(q)}{p_{m,1}(q)}s}{p_1(q)/p_{m,1}(q)},$$

$$p_{s0}/p_{m0} = a_{m0}/a_{s0},$$

$$p_{s1}(q)/p_{m1}(q) = a_{m1}(q)/a_{s1}(q),$$

$$p_0/p_{m,0} = \left[\omega_{c,m} + \omega_{c,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}},$$

$$p_{k,0}/p_{m,0} = \left[\omega_{i,m} + \omega_{i,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}},$$

$$p_0/p_{k,0} = \frac{\left[\omega_{c,m} + \omega_{c,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}}{\left[\omega_{i,m} + \omega_{i,s}\left(\frac{p_{s,0}}{p_{m,0}}\right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}},$$

$$p_1(q)/p_{m,1}(q) = \left\{ \omega_m + \omega_s\left[\frac{p_{s,1}(q)}{p_{m,1}(q)}\right]^{1-\varepsilon} \right\}^{\frac{1}{1-\varepsilon}}.$$

Equation (13) can be solved numerically for the amount of savings in period 0, k_1. In Figure 5 we consider the same parametrization used in the benchmark economy in the text (see Table 9), but we now allow for different levels of risk aversion. We can observe that increasing (decreasing) the degree of risk aversion, the effects of volatility on savings, consumption expenditure and the share of services in both consumption and value added increase (decrease).

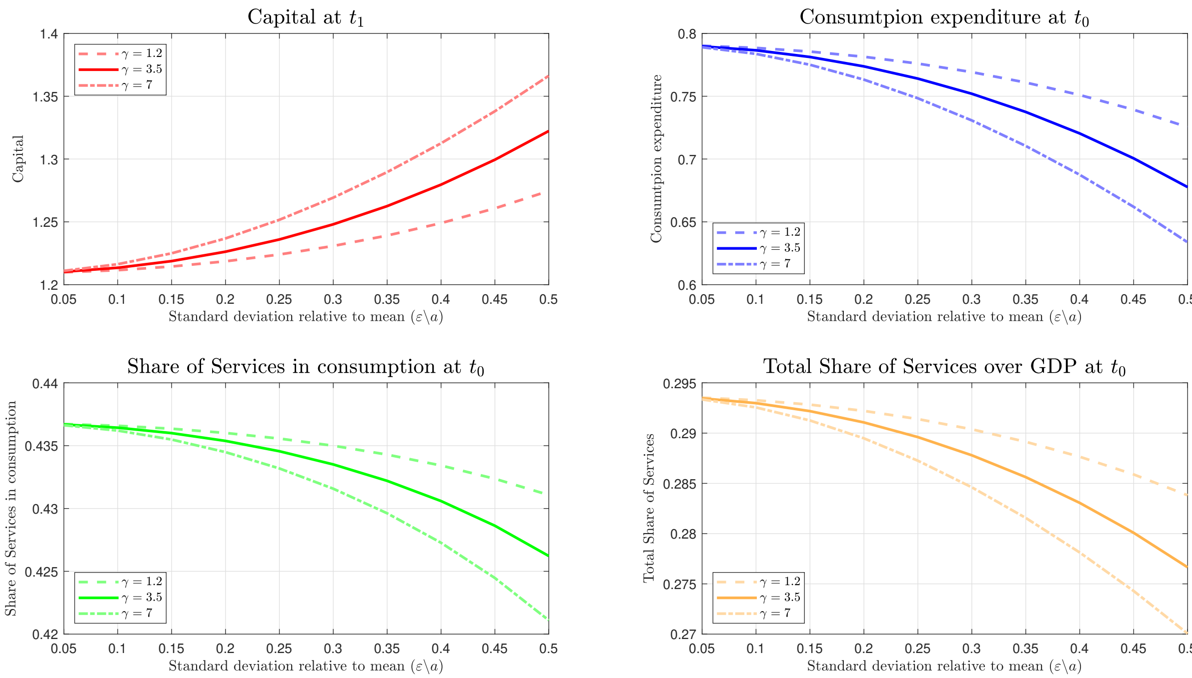

Figure 5: Different risk adversion

Table 9: Model parameters

Parameters	**Symbol**	**Value**
Intertemporal elasticity of substitution	ψ	$\frac{1}{3.5}$
Discount factor	β	0.95
Minimum required consumption in services	s	0.1
Elasticity of substitution between manufacturing and services	ε	0.5
Initial capital	k_0	1
Stone-Geary consumption weights	$\omega_{c,m} = \omega_{c,s}$	0.5
Stone-Geary investment weight – manufacturing	$\omega_{i,m}$	0.8
Stone-Geary investment weight – services	$\omega_{i,s}$	0.2
Total factor productivities (at $t = 0$)	$a_{m,0} = a_{s,0}$	2
Probabilities of each states	$P_1 = P_2 = P_3 = P_4$	0.25