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RESEARCH ARTICLE

Cyclical labour income risk in Great Britain

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Summary

This paper provides new evidence on the cyclical behaviour of household labour income risk in Great Britain and the role of social insurance policy in mitigating against this source of income risk. To achieve this, we decompose stochastic idiosyncratic household income into its transitory and persistent components. We focus our analysis of income risk captured by the second to fourth moments of the probability distribution of shocks to the persistent component of income. We find that household labour income risk increases during contractions via changes in third and fourth central moments of persistent shocks to labour income, whereas the variance remains acyclical. We also find that economic policy has reduced the level of risk exposure and its increase during contractions via benefits rather than tax policies.

KEYWORDS

aggregate fluctuations, household income risk, social insurance policy

1 | INTRODUCTION

The cyclical behaviour of labour income risk and the extent to which risk exposure increases during periods of contraction have significant implications for household welfare and economic policy. The effects of adverse idiosyncratic labour income shocks on households' consumption fluctuations are typically stronger for persistent relative to purely transitory shocks.[1] Therefore, knowing whether periods of contraction imply an increased probability of receiving large persistent negative labour income shocks is important when examining the cyclicality of labour income risk. The absence of market opportunities for complete insurance against negative shocks to labour income motivates public intervention. This intervention involves providing insurance typically via the redistributive mechanisms embedded in the tax system and the insurance element of benefits policies. Thus, understanding the cyclical properties of labour income risk and which policies successfully mitigate the increase in risk exposure during contractions is critical for policymaking.

These considerations have motivated empirical research which examines whether the moments of the distribution of shocks to labour income higher than the first depend on the aggregate state of the economy. Given the importance of persistent labour income shocks, a small number of studies have directly examined the cyclical properties of the shocks to the persistent component of the individual or household income. In a seminal analysis, Storesletten et al. (2004) estimated a model for income dynamics with a state-dependent variance, using US survey data from the Panel Study of Income Dynamics (PSID), and found that the variance of the persistent component of household labour income is countercyclical. The advantage of their methodology is that it identifies the contrast between booms and slumps by exploiting the history dependence of cross-sectional moments, which incorporate cyclical variation at the aggregate level during and before the panel data sample period. In an influential contribution, Guvenen et al. (2014) study the distribution of individual males

[1]See, for example, Meghir and Pistaferri (2011) for a review of research on earnings dynamics.

earnings growth using US Social Security Administration data. They discover that its left-skewness is countercyclical, whereas its variance is acyclical. These imply an increased probability of negative shocks to labour income during contractions. Busch and Ludwig (2016, 2020), using data for Germany and the United States, extend the approach in Storesletten et al. (2004) and estimate models for income dynamics that allow for regime-switching second, third and fourth central moments. They confirm that the moments of shocks to the persistent component of household labour income higher than the first are cyclical, implying an increase in risk during contractions. They also find that government intervention, in the form of taxation and social insurance/benefits policies reduces the increase in risk exposure arising from shocks to permanent labour income in contractions. However, they do not examine the effect of the tax and benefit element separately.

We provide new evidence on the cyclical behaviour of household labour income risk in Great Britain (GB) and assess the roles of tax and benefit policy in mitigating against this source of risk. We achieve this using panel data for 1991–2008 from the British Household Panel Survey (BHPS) and aggregate time series of the economic cycle since 1956.[2] To capture fluctuations in the aggregate state so that we can study its implications for idiosyncratic risk, we classify years as expansions or contractions, based on the official classification of years as recessions following the Bank of England definition. This requires two consecutive quarters of contraction in GDP. The BHPS dataset, which has been used extensively for income dynamics analysis in the United Kingdom, provides measures of annual earnings at the individual and household levels, in addition to observable characteristics.[3] We examine the effect of economic policy using different measures of pre- and post-policy household income.

We apply a model of income dynamics that decomposes stochastic idiosyncratic household income into components that capture initial (cohort-specific) conditions as well as transitory and persistent shocks. We let the second through fourth central moments of the probability distributions of these shocks depend on calendar time. In particular, following the parametric approach of Storesletten et al. (2004) and Busch and Ludwig (2016, 2020), we allow the moments of transitory and persistent shock distributions to vary between expansions and contractions of the aggregate economy. Motivated by empirical evidence for GB, we also allow the moments of the cohort-specific distributions to be time-varying to capture underlying cohort-level heterogeneity that may relate to pre-labour market conditions, separately from cyclical fluctuations. Our model is estimated employing a minimum-distance estimation procedure.

Our first set of findings relates to the cyclicality of risk in household labour income. We find that risk increases during contractions in GB and that this is due to changes in the third and fourth central moments of the distribution of shocks to the persistent component of labour income. The more robust effect comes from the third central moment, which becomes more negative during contractions. In contrast, we do not find evidence of cyclicality for the variance.[4] Note that a negative third central moment signifies that the left tail is thicker than the right tail. A thicker left tail in contractions than in expansions implies a higher probability of a household receiving a large negative persistent income shock in bad times. Moreover, an increase in the fourth central moment during contractions of the aggregate state works to increase the probability of receiving extreme shocks, implying an even thicker left tail. Our results are broadly consistent with the findings reported in Guvenen et al. (2014) and Busch and Ludwig (2020) for the United States, Busch and Ludwig (2016) for Germany and Busch et al. (2021) for Germany, Sweden, France and the United States.

Our second set of results relates to evaluating the risk mitigation effects of taxation separately from benefits. We find that social insurance works in this respect via benefits. These significantly reduce both the level of risk exposure and its increase during contractions, when focusing on persistent shocks to household labour income. In contrast, taxes do not have a significant effect. The separate evaluation of taxes and benefits to reduce the cyclicality of the persistent component of labour income risk is novel in this literature. Looking at earnings growth as a proxy for earnings risk, Busch et al. (2021) find that both taxes and benefits reduce the cyclicality in risk exposure associated with changes in the skewness of the distribution of earnings growth.[5]

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[2]The longer time series for the aggregate state is used because the moments of the persistent component of the idiosyncratic shocks incorporate past cyclical variation since 1956. This period corresponds to the year that 60-year-old individuals in 1991 entered the labour market at age 25.

[3]Household characteristics allow us to partial out observable deterministic components (i.e. experience, education, region of residence and household size effects) to isolate idiosyncratic labour income in the data.

[4]For GB, Bayer and Juessen (2012) follow the approach in Storesletten et al. (2004) and find that the variance of idiosyncratic shocks to wages is acyclical using the BHPS dataset. Moreover, Cappellari and Jenkins (2014), using BHPS data, find that the variance of individual earnings growth shows little time variation.

[5]Evidence from, for example, Blundell and Etheridge (2010) and Belfield et al. (2017) demonstrates that the UK benefits have stronger effects than taxes in mitigating household income inequality. Here, we focus on the effects of social insurance on risk reduction.

We organise the rest of the paper as follows. In the next section, we provide an overview of the data and relevant properties that motivate the model specification. In Section 3, we introduce the theoretical model of income risk dynamics that we estimate. The results are in Sections 4 and 5. Finally, Section 6 contains the conclusions.

2 | DATA AND EMPIRICAL MOTIVATION

We next provide information on the BHPS dataset and variables used for the estimation and analysis of the empirical properties of inequality that underlie our model specification and identification in Section 3.

2.1 | Data

The BHPS is a comprehensive longitudinal survey for GB, covering 1991 to 2008.[6] This dataset includes information for up to 5000 households on earnings and other sources of income for individuals and households over an annual period starting in September and on the socio-economic characteristics of the respondents. These characteristics include gender, education, age, social (professional) class and region.[7] BHPS was replaced in 2010 by a new panel data survey, Understanding Society, which does not include information on annual earnings, and thus cannot be used to analyse annual earnings risk. We also use the auxiliary dataset Derived Current and Annual Net Household Income Variables (DCANHIV), compiled by Bardasi et al. (2012), which contains derived data on household disposable income. Note that the DCANHIV dataset tracks the same individuals/households for the same period as the BHPS, that is, 1991–2008.

We retain all households where the head is between 23 and 62 years for our analysis. The household heads must have non-imputed individual earnings and report individual earnings higher than half the product of the minimum legal hourly wage and 520 h.[8] We further restrict the sample by keeping the households who are present for at least three consecutive periods.

The DCANHIV dataset provides consistent series for household labour income, private transfers, taxes and national insurance contributions and benefits. Labour income is the sum of annual earnings of the household members plus annual private transfers income. Taxes refer to annual household income taxes after credits, whereas benefits are the annual social benefits income, which totals all receipts from state benefits from all household's members (including national insurance retirement pensions). We present in our analysis below results for labour income, labour income minus taxes and national insurance contributions, labour income plus benefits, and labour income net of taxes and national insurance contributions plus benefits.

2.1.1 | Aggregate state

Given that the BHPS annual cycle begins in September, a year t in our analysis refers to the date September $t-1$ to August t. We classify the years between 1956 and 2008 into recessions and expansions depending on whether they include an official recession. We define recessions following the Bank of England practice, which requires two consecutive quarters of contraction in GDP. Following this, a year t in our analysis is a year of recession when $t = 1956, 1961, 1962, 1973, 1974, 1975, 1980, 1981, 1990, 1991, 2008$.

2.2 | Inequality over cohorts

We next examine the cohort and age-specific properties of the cross-sectional distribution of residual household labour income, net of variation due to observable characteristics. The illustration in Figure 1 reveals empirical properties used to motivate our model specification. Following Deaton and Paxson(1994) (see also Storesletten et al. (2004), Heathcote et al. (2005) and Blundell et al. (2015)), we use a dummy variable regression decomposition of the

[6]Further information on the datasets, definitions and construction of variables and details on sample selection, can be found in Appendix A of the supporting information.

[7]Data on Northern Ireland are available from 1997 via the additional BHPS sub-sample European Community Household Panel Survey. However, we focus on GB to not further restrict the time dimension, which is important for our analysis.

[8]We follow Blundell and Etheridge (2010) for the definition of the head. The head is usually the male in a household consisting of a married couple with children or the oldest working male. See Appendix A.2 in the supporting information for details.

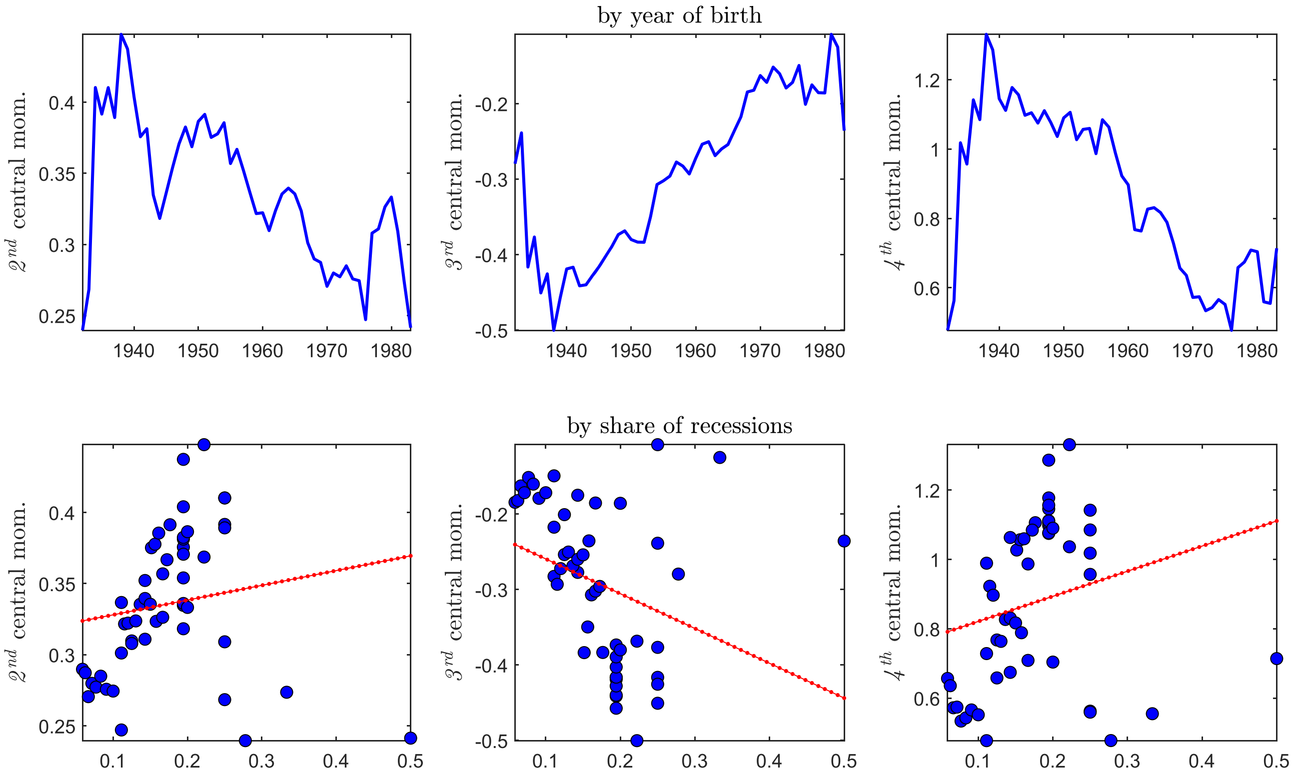

FIGURE 1 Cohort effects on the distribution of residual household labour income. Note: Figure 1 plots the coefficients of cohort-specific dummies in a regression of central moments of the cross-sectional distributions of residual household labour income on cohort and age dummies, for different year of birth (1st row) and different shares of recession years since the age of 25 (2nd row)

cross-sectional higher central moments of the idiosyncratic component of household labour income, to estimate cohort and age effects. The estimated idiosyncratic component, denoted as $\hat{v}_{i,h,t}$, for household i of age group h at time t, is obtained as the residual of a Mincerian type regression that partials out observable variation in household labour income.[9] Then, for every year t, we group households into 5-year adjacent age cells indexed by h, that is, we define a household as belonging to the age group h if the age of the head was between $h-2$ and $h+2$. We obtain the cross-sectional central moments $m_\tau^v(h,t,g)$, for $\tau = 2, 3, 4$, as

$$m_\tau^v(h,t,g) = \frac{1}{N_{h,t}} \sum_{i=1}^{N_{h,t}} \left[\left(\hat{v}_{i,h,t} - \bar{\hat{v}}_{i,h,t} \right)^\tau \right], \tag{1}$$

where $g = t-(h+24-1)$ denotes the cohort, $\bar{\hat{v}}_{i,h,t}$ denotes the sample averages of $\hat{v}_{i,h,t}$, $N_{h,t}$ is the number of households i of age group h at time t.[10] We then regress $m_\tau^v(h,t,g)$'s on a full set of age and cohort dummies, that is,

$$m_\tau^v(h,t,g) = \beta_{h,\tau}' D_h + \beta_{g,\tau}' D_g + \nu_{h,t,g,\tau},\ \tau = 2, 3, 4, \tag{2}$$

where $\nu_{h,t,g,\tau}$ is an error term.[11] Figure 1 presents different plots using the estimated $\beta_{g,\tau}$, capturing cohort-related effects on the distribution of residual household labour income in relation to initial conditions (birth year) and experience of recessions, partialling out age effects.

The first row of Figure 1 depicts the cohort effects (i.e. the estimated cohort coefficients) over the period that defines the cohort (i.e. year of birth). These plots show a downward trend in cohort second and fourth central moments for newer cohorts and an upward trend in the third central moment. These trends imply that the cohort effects are drawn

[9]The Mincerian regression will be discussed in more detail in the next section.

[10]This implies a total of $3 \times T \times H = 1{,}944$ moments $m_k^v(h,t,g)$, because our sample length and age groupings are $T = 18$ and $H = 36$, respectively.

[11]Note that we cannot control for cohort, age and time effects simultaneously due to perfect collinearity.

from distributions that tend to become more symmetric over time and with less extreme values, further suggesting that controlling for life-cycle effects, younger generations are more similar in terms of household labour income variation. This observation may reflect fewer or milder aggregate macroeconomic shocks for younger generations. However, it may also capture more similarity in labour market skills at the point of entry into the labour market, which may arise from social changes and government intervention (e.g. public education and health). These observations help to motivate our modelling of effects associated with initial conditions that remain over the work-life. In particular, we will partial out the influence of potential pre-labour market changes on labour income risk to focus on the importance of shocks during work-life.

The second row, where we plot the cohort coefficient estimates against the proportion of contractions that each cohort has experienced, is motivated by Figure 1b in Storesletten et al. (2004). Using PSID data in the United States, Storesletten et al. (2004) find that cohort variance is increasing in the share of recessions that the cohort has experienced. This finding motivates their interest in estimating the aggregate-state dependency of the variance by exploiting between-cohort variation in their experience of recessions. Row 2 in Figure 1 suggests that for GB the variance seems unrelated to the share of recessions. In contrast, there is some evidence of a negative relationship between the third central moment and the share of recessions and a positive relationship with the fourth central moment. As pointed out above, because the estimated cohort coefficients do not control for time, the relationships shown in row 2 may be affected by generation-specific pre-labour market factors. Thus, it is important in our model of the idiosyncratic component of income developed in the next Section to allow for generation-specific initial conditions in addition to dependence on the aggregate state.

3 | INCOME RISK MODEL

Our econometric analysis is motivated by the framework and identification strategy introduced by Storesletten et al. (2004) and extended by Busch and Ludwig (2016, 2020) to study persistent shocks that vary with the aggregate state. This analysis builds on a large literature of income dynamics (see Topel and Ward (1992), Gottschalk and Moffitt (1994), Haider (2001), Baker and Solon (2003), Kalwij and Alessie (2007), Blundell et al. (2015) and Meghir and Pistaferri (2011) for a review). The basic object of analysis for the various measures of income and risk is households whose head is aged between 25 to 60 in the period 1991–2008.

3.1 | Model specification

We assume that the process determining the natural logarithm of annual household income, $y_{i,h,t}$, is composed of an observable deterministic part, $q\left(x_{i,h,t}\right)$, and an unobservable random component, $\upsilon_{i,h,t}$:

$$y_{i,h,t} = q\left(x_{i,h,t}\right) + \upsilon_{i,h,t}, \tag{3}$$

for household i whose head has age $h = 1, \ldots, H = 36$ in period $t = 1991, \ldots, 2008$.[12] The idiosyncratic component of income, $\upsilon_{i,h,t}$, is driven by: (i) stochastic effects that are received once at birth, remain fixed over the lifetime and that depend on the cohort, $\chi_{i,g}$, where $g = t-(h+24-1) \in [1932,1933, \ldots, 1984]$ denotes year of birth (i.e. the cohort effect); (ii) persistent shocks, $z_{i,h,t}$, and (iii) transitory shocks, $\varepsilon_{i,t}$. In particular, $\upsilon_{i,h,t}$ and $z_{i,h,t}$ are given by

$$\upsilon_{i,h,t} = \chi_{i,g} + z_{i,h,t} + \varepsilon_{i,t}, \tag{4}$$

$$z_{i,h,t} = \rho z_{i,h-1,t-1} + \eta_{i,t}, \tag{5}$$

where $0 < \rho < 1$ and $\eta_{i,t}$ captures innovations to the persistent effects. The distributional assumptions for the three sub-components in (4) and (5) are given by

$$\chi_{i,g} \underset{i.i.d.}{\sim} F_{\chi}(0, m_2^{\chi} + b_2^{\chi}\gamma_g, m_3^{\chi} + b_3^{\chi}\gamma_g, m_4^{\chi} + b_4^{\chi}\gamma_g), \tag{6}$$

[12]Note that we define $h = 1$ for a 25 year old.

$$\varepsilon_{i,t} \underset{i.i.d.}{\sim} F_\varepsilon(0, m_2^{\varepsilon,f(t)}, m_3^{\varepsilon,f(t)}, m_4^{\varepsilon,f(t)}), \tag{7}$$

$$\eta_{i,t} \underset{i.i.d.}{\sim} F_\eta\left(0, m_2^{\eta,f(t)}, m_3^{\eta,f(t)}, m_4^{\eta,f(t)}\right), \tag{8}$$

where F_χ, F_ε, and F_η denote the density functions of $\chi_{i,g}$, $\varepsilon_{i,t}$ and $\eta_{i,t}$ respectively. The four arguments in each of these functions refer to the first through the fourth central moment. Moreover, $\gamma_g = g - 1931 \in [1, 2, \ldots, 53]$ is a cohort-defined trend tracking the birth years of the households whose head was aged between 25 to 60 in 1991-2008.

The distributions of the innovations to the transitory and persistent shocks are allowed to be time dependent. In particular, we let the respective moments higher than the first take values depending on whether the aggregate state, $f(t)$, refers to an expansion, e, or a contraction, c. We define two indicator variables: (i) $I_{f(t)=e}$, equal to 1 if period t is an expansion (and zero otherwise); and (ii) $I_{f(t)=c}$, equal to 1 if period t is a contraction (and zero otherwise). This implies that for $k = \varepsilon, \eta$ and $\tau = 2, 3, 4$:

$$m_\tau^{k,f(t)} \equiv m_\tau^{k,e} I_{f(t)=e} + m_\tau^{k,c} I_{f(t)=c}. \tag{9}$$

The model specification implies that the history of persistent shocks after entering the labour market at the age $h = 1$ is important for idiosyncratic income. We assume that $z_{i,0,t} = 0$, implying that before joining the labour market there are no persistent shocks that matter for income dynamics after $h = 1$ other than those captured by the initial conditions $\chi_{i,g}$. In turn, these remain fixed over the work-life and thus do not have a time subscript. In other words, the initial conditions represent stochastic factors that are relevant for income dynamics before joining the labour market. Motivated by empirical observations in Figure 1, we have allowed for the possibility that the effects $\chi_{i,g}$ differ across generations. Given that our interest is in estimating potential dependence of $m_\tau^{k,f(t)}$, $k = \varepsilon, \eta$ and $\tau = 2, 3, 4$, on contractions, we let the initial conditions vary over cohorts to not confound the time-dependence of working life idiosyncratic shocks with underlying time variation of pre-work heterogeneity.

The model specification here is a generalisation of those in Storesletten et al. (2004) and Busch and Ludwig (2016, 2020). In particular, if we set $b_2^\chi = b_3^\chi = b_4^\chi = 0$, $(m_2^{\varepsilon,f(t)}, m_3^{\varepsilon,f(t)}, m_4^{\varepsilon,f(t)}) = (m_2^\varepsilon, m_3^\varepsilon, m_4^\varepsilon)$, the model is the same with that in Busch and Ludwig (2020). If we further also assume that $m_3^\varepsilon = 0$, that $m_3^{\eta,e} = m_3^{\eta,c} = 0$ and $m_4^{\eta,e} = 3\left(m_2^{\eta,e}\right)^2$, and $m_4^{\eta,c} = 3\left(m_2^{\eta,c}\right)^2$, we return to the base specification in Storesletten et al. (2004).[13]

The set of theoretical moments for $v_{i,h,t}$ are a function of past moments of innovations to the persistent component (see Appendix B.1 in the supporting information).[14] Therefore, estimation of the parameters of interest requires knowledge of whether $h-1$ years before those in the observed sample of households were expansionary or contractionary. In turn, this implies that more time variation in the aggregate state is exploited in the estimation. Thus, helping to increase the accuracy of estimating moments separately for periods of expansion and contraction (see also Storesletten et al. (2004), who introduced this identification approach and the web appendix of Bayer and Juessen (2012) for a Monte Carlo evaluation of this method).

3.2 | Estimation

We next calculate the empirical central moments of the residuals from the Mincer regression. To obtain the idiosyncratic component of household income, $v_{i,h,t}$, we follow the literature on earnings dynamics and run a Mincerian-type regression to partial out non-stochastic effects from income. In particular, we assume that (3) is given by

$$y_{i,h,t} = bx_{i,h,t} + v_{i,h,t}, \tag{10}$$

where b is a vector of parameters.[15] The regressors, $x_{i,h,t}$ in (10) include calendar year time effects, dummies for experience (captured by a full set of age dummies), region of residence dummies, gender dummies, marital status dummies,

[13] In this literature, as well as here, the time series of the aggregate state is taken as given. It is an interesting and non-trivial extension in this framework to model the aggregate state jointly with idiosyncratic shocks (see Bloom et al. (2018) on the importance of firm-level uncertainty for aggregate fluctuations).

[14] In Appendix E of the supporting information, we discuss an alternative representation of co-kurtosis and summarise the main results under this specification to establish robustness.

[15] For the estimation, we normalise $t = 1, \ldots, T = 18$.

household size and interaction between marital status dummies and household size. For the region dummies, we use the UK Government Office Regions classification which corresponds with the highest tier of sub-national division in England, Scotland and Wales. Furthermore, following Meghir and Pistaferri (2004), we allow for the returns to the observable deterministic characteristics to be skill-specific. Hence, we estimate (10) for two separate skill groups, that is, households whose head has a University education and those households whose head does not. Finally, because in our econometric analysis, we employ household quantities for the arguments in (10), we define the age, gender, marital status and regional effects in terms of the head of the household. We denote by $\hat{v}_{i,h,t}$ the estimated idiosyncratic component of household income.

In every year t, we group households in the sample into 5-year adjacent age cells indexed by h, that is, we define a household as belonging to the age group h if the age of the head was between $h-2$ and $h+2$. For example, the first cell, that is, age group 25, contains households with age between 23 and 27 years old, the second cell, that is, age group 26, contains all households with heads between 24 and 28 years old, whereas the last cell, that is, age group 60, contains all households with heads between 58 and 62 years old. The empirical moments are given by:

$$\frac{1}{I_{h,t,\kappa}} \sum_{i=1}^{N} \iota_{i,h,t,\kappa} \left[\left(\hat{v}_{i,h,t} - \bar{\hat{v}}_{i,h,t} \right)^{\phi} \left(\hat{v}_{i,h+\kappa,t+\kappa} - \bar{\hat{v}}_{i,h+\kappa,t+\kappa} \right)^{\psi} \right], \tag{11}$$

where $\bar{\hat{v}}_{i,h,t}$ and $\bar{\hat{v}}_{i,h+\kappa,t+\kappa}$ denote the sample averages of $\hat{v}_{i,h,t}$ and $\hat{v}_{i,h+\kappa,t+\kappa}$ respectively, $(\phi, \psi) \in \{(1,1),(2,1),(2,2)\}$, $\kappa = 0, \ldots, \min\left[T-t, H-h\right]$, $I_{h,t,\kappa} = \sum_{i=1}^{N} \iota_{i,h,t,\kappa}$ and ι is an indicator function which is one when an individual i of age group h at time t is also present in time $t+\kappa$, and zero otherwise. This implies a total of $3 \times \sum_{t=1}^{T} \sum_{h=1}^{H} \min\{H-h+1, T-t+1\}$ moments. Our sample length and age grouping of $T = 18$ and $H = 36$ thus lead to a total of 15,561 empirical moments that we calculate from the data. In contrast, our theoretical moments are a function of the 19 parameters in the theoretical model, in which the moments higher than the first vary across age and time following the parametric restrictions specified by the model. Conditional on specific sequences of contractions and expansions of the aggregate state $f(t)$ observed in the data, we estimate the parameters employing a minimum distance estimator that chooses the parameter vector to minimise the distance between the empirical and the theoretical moments. We discuss the details on estimation and the bootstrap procedure used to estimate the confidence intervals for hypothesis testing in Appendix B.2 of the supporting information. In Appendix B.3 of the supporting information, we discuss how variation in the aggregate state informs the estimation. Finally, as shown in Appendix B.4 of the supporting information, given sufficient variation in the history of the aggregate state, the 19 parameters in the theoretical model are identified if we observe four consecutive periods and four age groups in the panel sample.

4 | TIME VARIATION IN LABOUR INCOME RISK

We start with an analysis of the main result relating to the cyclicality of the persistent process and then discuss extensions.

4.1 | Higher-order cyclicality of persistent shocks

We summarise the estimated model parameters relating to the persistent component of labour income shocks in Table 1. The estimates reveal that the third central moments of the distribution of shocks to the persistent component of idiosyncratic labour income (i.e. of the distribution F_η in Section 3) are negative in both expansions and contractions. These results indicate negatively skewed distributions of persistent shocks, implying a higher probability of receiving large negative instead of positive shocks. Moreover, the results in Table 1 show that the fourth central moment in recessions is higher than that of a normal distribution.

To further examine whether the central moments of F_η are higher in absolute value during contractions, we report in Table 2 the difference in these moments between expansions and contractions. Table 2 also shows results from a one-sided test of the null hypothesis that the moments in expansions are higher than or equal to (in absolute value) the respective moments in contractions. Table 2 reveals that, whereas the variance of the shocks is not significantly higher in contractions relative to expansions, the third central moment is significantly more negative. The fourth central moment is significantly higher at the 10% level when comparing official recessions with expansions.

The cyclical pattern of the third and fourth central moments imply cyclical variation in risk. First, conditional on the fourth central moment, the asymmetry in the distribution of idiosyncratic labour income shocks, implied by the negative

TABLE 1 Persistent labour income process

[1] ρ	[2] $m_2^{\eta,e}$	[3] $m_2^{\eta,c}$	[4] $m_3^{\eta,e}$	[5] $m_3^{\eta,c}$	[6] $m_4^{\eta,e}$	[7] $m_4^{\eta,c}$
0.7391***	0.0654***	0.0701***	−0.0241 **	−0.1025***	0.0149	0.1549*

Note: The H_0's $\rho = 0$, $m_2^{\eta,j} = 0$, $m_3^{\eta,j} = 0$ for $j \in (\{c\}, \{e\})$ are rejected at significance level 1% (***), 5% (**) or 10% (*), based on a two-tail confidence interval procedure. For m_4, the H_0 is $m_4^{\eta,j} = 3\left(m_2^{\eta,j}\right)^2$, where $j \in (\{c\}, \{e\})$ and H_A is $m_4^{\eta,e} \geq 3\left(m_2^{\eta,j}\right)^2$, and is rejected at the same significance levels based on a one-tailed confidence interval. All the tests are implemented using a block bootstrap with 3000 replications.

TABLE 2 Tests of procyclicality

[1] $m_2^{\eta,e} - m_2^{\eta,c}$	[2] $m_3^{\eta,e} - m_3^{\eta,c}$	[3] $m_4^{\eta,e} - m_4^{\eta,c}$
−0.0047	0.0784***	−0.1400*

Note: The quantities in the table have been calculated by using the estimates in Table 1. The H_0's $m_2^{\eta,e} - m_2^{\eta,c} \geq 0$, $m_3^{\eta,e} - m_3^{\eta,c} \leq 0$, $m_4^{\eta,e} - m_4^{\eta,c} \geq 0$ are rejected at significance level 1% (***), 5% (**) or 10% (*), based on a one-tailed confidence interval procedure, implemented using a block bootstrap with 3000 replications.

third central moment in contractions, suggests that in downturns it is more likely to draw large negative relative to large positive labour income shocks. A negative third central moment indicates that the left tail is thicker than the right tail. Because the left tail represents negative shocks, a thicker left tail in contractions than in expansions implies a higher probability of a household receiving a very large negative income shock in bad times.[16] Second, conditional on the third central moment, an increase in the fourth central moment implies an increase in the probability of receiving extreme shocks. Therefore, because the fourth central moment is not reduced in recessions, the probability of receiving extreme shocks is at least as high in these periods. Together, the third and fourth central moments imply an increase in labour income risk in downturns. Therefore, overall we have evidence of countercyclical risk in Great Britain.

The cyclicality of higher-order household labour income risk for GB coheres with international evidence in Busch and Ludwig (2016, 2020) who decompose shocks to household labour income. They also investigate the cyclicality of third and fourth central moments of shocks to the persistent component separately from possible initial conditions and transitory shocks. Despite the differences in our modelling of the income process, discussed in Section 3, our findings for GB are similar to those in Busch and Ludwig (2016, 2020) for Germany and the United States. In addition to cyclical third and fourth moments, these studies also find that the variance of shocks to persistent income increases during contractions. In contrast, our findings for GB highlight the importance of the asymmetry of the distribution of shocks without significant cyclical changes in the variance. The significance of the cyclicality of the asymmetry of the distribution of shocks is also consistent with the findings of Guvenen et al. (2014) and Busch et al. (2021). They emphasise the countercyclical property of the left-skewness of income risk when the latter is approximated by earnings growth.[17] The results here broadly cohere with those of Bayer and Juessen (2012), who also find an acyclical variance of wage risk for the United Kingdom.[18] These findings are also generally consistent with evidence for GB reported in Blundell and Etheridge (2010) who decompose household earnings shocks into permanent and transitory components. The estimated variances of both earnings shock components over 1991-2003 in Blundell and Etheridge (2010) do not show evident co-movement with the aggregate conditions.

The results in Table 2 demonstrate cyclical risk in terms of shocks to the persistent component of labour income. In contrast, the results regarding the distribution of shocks to the transitory component of labour income show that, despite

[16]Guvenen et al. (2014) refer to this change in the distribution as countercyclical left-skewness. This is consistent with the results here, that is, in a contraction, the third central moment is smaller (i.e. more negative) than in an expansion.

[17]Guvenen et al. (2014) and Busch et al. (2021) approximate income risk by income growth, without a statistical decomposition of shocks to those affecting the persistent and the transitory component of income. However, Guvenen et al. (2014) also examine the moments of the distribution of 5-year earnings growth to approximate more persistent shocks.

[18]Using PSID data for the United States, Storesletten et al. (2004) find a countercyclical variance of shocks to the persistent component of household labour income. Ziliak et al. (2011), employing data from the Current Population Survey, without decomposing shocks to transitory and persistent components, found that the volatility of individual male and female earnings growth are countercyclical and procyclical, respectively.

TABLE 3 Probabilities of big shocks

		[1] $\eta_{i,t} \sim \hat{F}_\eta$	[2]	[3] $\eta_{i,t} \sim N\left(0, m_2^{\eta, f(t)}\right)$	[4]
Probability of		**Recession**	**Expansions**	**Recessions**	**Expansions**
income gain > 40%:	$Pr(\eta_{i,t} > \ln(1.40))$	0.00%	0.00%	10.28%	9.52%
income gain > 20%:	$Pr(\eta_{i,t} > \ln(1.20))$	0.00%	2.90%	24.65%	23.91%
income loss > 50% :	$Pr(\eta_{i,t} < \ln(0.50))$	3.15%	1.11%	0.45%	0.35%
income loss > 75%:	$Pr(\eta_{i,t} < \ln(0.75))$	2.07%	0.00%	0.00%	0.00%

Note: The probabilities of gains/losses are with respect to the level of persistent income under the mean shock. $\eta_{i,t}$ refers to the natural logarithm of persistent shocks to labour income and has zero mean.

differences in the distributions between expansions and contractions, the change is not readily associated with an increase in transitory risk in contractions (see Appendix Table C.1 in the supporting information).

The parameter estimates relating to the stochastic effects, captured by the initial conditions, are reported in Appendix Table C.2 of the supporting information. The results confirm the downward trend in variance suggested by the first row of Figure 1. However, once the time-variation in household income risk driven by aggregate conditions is accounted for, the time pattern of the third and fourth central moments of the distribution of pre-labour market shocks is not statistically significant.

4.1.1 | Economic intuition

To illustrate the income risk implications of the differences in the third and fourth central moments of F_η between recessions and expansions, we next calculate probabilities of labour income shocks associated with the tails of the distribution under two different assumptions.[19] First, we assume that F_η is an asymmetric distribution, which we calibrate to the second to fourth moments estimated from the data. Second, we assume that this distribution is instead given by a symmetric normal distribution. In this case, only the variance changes between recessions and expansions.

To operationalise this approach, we compute a distribution F_η following the approach in Guvenen et al. (2014) by assuming that this distribution is a mixture of two Gaussian distributions in expansions, and a different mixture in recessions. In particular, we assume that

$$\eta_{i,t} \sim \hat{F}_\eta = \begin{cases} \eta^1_{i,t} \sim N(\mu_{1,f(t),}, \sigma_1^2) & \text{with probability } p_{f(t)} \\ \eta^2_{i,t} \sim N(\mu_{2,f(t),}, \sigma_2^2) & \text{with probability } 1 - p_{f(t)} \end{cases}$$

where $0 < \sigma_1^2, \sigma_2^2 < \infty$ and $0 \le p_{f(t)} \le 1$. We calibrate the parameters $\{p_e, p_c, \mu_{1e}, \mu_{2e}, \mu_{1c}, \mu_{2c}, \sigma_1^2, \sigma_2^2\}$, using a minimum distance routine, so that the calibrated distribution, $\hat{F}_\eta$, has the same central moments as F_η when the aggregate state is measured by the official classification in recessions. Using $\hat{F}_\eta$ we calculate the probability that a household receives big adverse and big positive shocks by integrating $\hat{F}_\eta$ below and above certain thresholds for expansions and recessions separately.

In Table 3, we report the probability that a household receives an idiosyncratic shock to the persistent component of their labour income that reduces (increases) their persistent income by a specific percentage, compared with what it would have been, had it received the mean $\eta_{i,t}$ shock (which is zero). To contextualise the shocks considered in Table 3, an example of a loss of 50% of annual labour income would be a situation where a member of a household with two working members becomes unemployed for the whole year, or both members face prolonged unemployment. In contrast, a loss of more than 75% implies further significant earnings losses for the other member as well. Further situations might include severe drops in earnings, for example, for self-employed members, or when shocks imply a change in a job associated with a big drop in wage when only one member works. The significant common element from these examples is that losses of 50% or 75% of household labour income reflect catastrophic events at the household level. In other words, they do not simply refer to an unemployment spell or a wage cut.

[19]More generally, the importance of the higher-order moments of income risk in explaining properties of earnings and wealth inequality in the data, and in evaluating the welfare implications of risk, has been analysed in the literature (see McKay (2017), De Nardi et al. (2020) and Angelopoulos et al. (2020)).

The results using $\hat{F}_\eta$ are in columns [1] and [2]. We can see that, during recessions, the left tail of the distribution becomes 'thicker'. In other words, the probability that a household receives a big adverse shock increases. At the same time, the right tail of the distribution becomes 'thinner'. That is, the probability a household receives a big beneficial shock decreases. Quantitatively, the effects regarding the left tail are important. In particular, the probability that a household loses more than half of its annual labour income more than doubles in recessions. Note that this probability is of a comparable magnitude with the proportion of long-term unemployed in the UK labour force. For example, between 1983-2008 this rate was, on average, 2.8% (see OECD, 2020).[20] Moreover, in more than 60% of the cases where households lose more than half of their income, households lose more than three-quarters of their labour income. This circumstance is in stark contrast with expansions, where even households who lose more than half of their labour income, do not lose more than three-quarters of it. In the context of the employment example, this implies that even if one member becomes unemployed for the whole year during expansions, the household can rely on the other member's employment to alleviate part of the shock. Therefore, catastrophic events, which have low probability and a huge impact, effectively do not happen in expansions. However, they do become possible in recessions.

To contextualise the effects resulting from the asymmetric distribution $\hat{F}_\eta$, we calculate the probabilities in Table 3 under a counterfactual experiment obtained by assuming that $\eta_{i,t}$'s are normally distributed with the same variance as for $\hat{F}_\eta$. We compute the distribution for the shocks to the persistent component of labour income as $\eta_{i,t} \sim N\left(0, m_2^{\eta,f(t)}\right)$. Two results stand out in this case. First, quantitatively, the increase in the probability of very bad shocks is very small, and catastrophic events remain zero probability events even in recessions. Second, qualitatively, the results speak of an increase in uncertainty during recessions, but not necessarily in downside risk, because the probability of big positive shocks also increases.

4.2 | Frequency of recessions

Recessions are not evenly distributed over the period 1956 to 2008. In particular, there are more frequent recessions until 1991, and no official recessions recorded after 1991, until the 2008 recession. We exploit this variation in the data to design an alternative test of the implications of cyclicality in labour income risk. The idea is that if recessions imply higher risk than expansions, then periods with a higher frequency of recessions should be associated with higher risk relative to periods with a lower frequency of recessions. More specifically, if shocks to the persistent component of residual labour income in recessions have higher third and fourth moments, but not second moments, relative to expansions, then periods with a higher frequency of recessions should be associated with higher income risk with the same form. In particular, with higher third and fourth moments, but not second moments, relative to periods with lower frequency of recessions.

To operationalise this experiment, we modify the model in (4)–(8), by testing whether labour income risk is higher in earlier periods relative to later periods. In particular, we let the respective moments higher than the first of the distribution in (8) take values depending on whether the aggregate state, $f(t)$, refers to the time period $A = [1956, \Lambda]$ or $B = [\Lambda+1{,}2008]$, where we set $\Lambda = 1991{,}1997$. This implies that for $\tau = 2, 3, 4$:

$$m_\tau^{\eta,f(t)} \equiv m_\tau^{\eta,e} I_{f(t)=A} + m_\tau^{\eta,c} I_{f(t)=B}. \tag{12}$$

Our motivation for choosing $\Lambda = 1991$ is that this year defines the start date for the longest time horizon since 1956 without official recessions, that is, the period 1991 to 2008 (which is fortunately within the BHPS sample). The beginning of the 1990s is also associated with a slow down in inequality (see Belfield et al. (2017)). The year $\Lambda = 1997$ further divides the sample into two periods with high and low frequency of recessions and is associated with a subsequent reduction in inequality which is also linked with a more robust macroeconomic performance (see Belfield et al. (2017)).

Estimation of the new model specification for income dynamics follows the same procedure as in the base model. In Table 4, we show the difference between the estimated moments for the two periods A and B, that is, for $\Lambda = 1991, 1997$.[21] The results confirm that there was an indeed higher risk in the earlier period of the sample, which has a higher proportion of recessions. In particular, Table 4 clearly shows that the higher risk in period A is associated with higher (in absolute

[20] Long-term unemployed include those who have been unemployed for at least 1 year.

[21] Note that the differences between periods reported in Table 4 are similar if Λ is set to different years in the 1990s.

TABLE 4 Differences between periods with different frequency of recessions

	$\Lambda = 1991$ Period A: 28% recessions Period B: 6% recessions	$\Lambda = 1997$ Period A: 24% recessions Period B: 9% recessions
$m_2^{\eta,A} - m_2^{\eta,B}$	0.0012	0.0017
$m_3^{\eta,A} - m_3^{\eta,B}$	−0.0528 **	−0.0467***
$m_4^{\eta,A} - m_4^{\eta,B}$	0.1660 **	0.0939***

Note: The H_0's $m_2^{\eta,A} - m_2^{\eta,B} \leq 0$, $m_3^{\eta,A} - m_3^{\eta,B} \geq 0$, $m_4^{\eta,A} - m_4^{\eta,B} \leq 0$ are rejected at significance level 1% (***), 5% (**) or 10% (*), based on a one-tail confidence interval procedure, implemented using a block bootstrap with 3000 replications.

TABLE 5 Policy effects on persistent component

	$\tau = 2$		$\tau = 3$		$\tau = 4$	
	$j = e$	$j = c$	$j = e$	$j = c$	$j = e$	$j = c$
$[m_\tau^{\eta,j}]^Y - [m_\tau^{\eta,j}]^{Y-T}$	0.0088	0.0127	−0.0067	−0.0265	0.0063	0.0509
$[m_\tau^{\eta,j}]^Y - [m_\tau^{\eta,j}]^{Y+bn}$	0.0168	0.0364**	−0.0233	−0.0925***	0.0125	0.1508**
$[m_\tau^{\eta,j}]^Y - [m_\tau^{\eta,j}]^{Y+bn-T}$	0.0246	0.0431***	−0.0239	−0.1001***	0.0132	0.1539**

Note: The superscripts Y, bn, T refer to labour income, benefits, and taxes and national insurance respectively. The H_0's $[m_2^{\eta,j}]^Y - [m_2^{\eta,j}]^\Gamma \leq 0$, $[m_3^{\eta,j}]^Y - [m_3^{\eta,j}]^\Gamma \geq 0$, $[m_4^{\eta,j}]^Y - [m_4^{\eta,j}]^\Gamma \leq 0$, for $\Gamma \in (\{Y-T\}, \{Y+bn\}, \{Y+bn-T\})$ and $j \in (\{c\}, \{e\})$, are rejected at significance level 1% (***), 5% (**) or 10% (*), based on a one-tail confidence interval procedure, implemented using a block bootstrap with 3000 replications.

value) moments $m_3^{\eta,f(t)}$ and $m_4^{\eta,f(t)}$, whereas differences in $m_2^{\eta,f(t)}$ are not significant, therefore being consistent with the explanation that the higher risk in period A is driven by the higher frequency of recessions because recessions imply exactly these characteristics of higher labour income risk. Our findings support the role of macroeconomic volatility as a contributor to increased household labour income risk before 1991; and vice versa, the role of greater macroeconomic stability as a contributor to the decline in household labour income risk post-1991.

Our results also contribute to the understanding of the evolution of income inequality in the United Kingdom. Previous research has established that the increase in household net income inequality in the United Kingdom, which had risen rapidly in the late 70s, slowed down after 1990 (see Blundell and Etheridge (2010) and Belfield et al. (2017)). Whereas household earnings inequality, and residual earnings inequality, increased until 1997, but at a slower rate relative to previous decades (see Blundell and Etheridge (2010)). Therefore, the period between 1990 and the late 1990s is a period of decompression in household earnings inequality, which was moderated even further after 1997. The moderation of inequality since the late 1990s has been linked to improved macroeconomic performance (Belfield et al. (2017)). Our results are consistent with these analyses of inequality, emphasising the role of macroeconomic stability in affecting household income risk and thus income differences between households.[22] In particular, our findings suggest that the more robust macroeconomic performance was working since the beginning of the 1990s to reduce household income risk, and thus positively impacted changes in income inequality.

5 | WHICH POLICY MITIGATES PERSISTENT SHOCKS?

We next estimate the model parameters using the different definitions of labour income post taxes and/or benefits, using the official recessions as the measure of the aggregate state and examine whether there is a decrease in risk post-policy.[23] In particular, we examine whether there is a significant reduction in the moments of the distribution of shocks to the persistent component of income, F_η, as a result of tax/benefit policy, by comparing moments estimated using labour income with those estimated using income post-policy intervention. To this end, Table 5 summarises the results of tests of the null hypothesis that the relevant moments are higher (in absolute value) post policy.

The differences between labour income and the various measures of income net of policy reported in Table 5 indicate, qualitatively, that tax and benefit policy together work in the right direction and reduce the variance and higher central

[22]For an analysis of income inequality in recent decades in the United Kingdom, see Jenkins (1995), Blundell and Etheridge (2010), Brewer and Wren-Lewis (2016), Belfield et al. (2017) and Blundell et al. (2018).

[23]The parameter estimates are reported in Appendix D of the supporting information.

TABLE 6 Policy effects on the cyclicality of persistent risk

	$\tau = 3$	$\tau = 4$
$[m_\tau^{\eta,e} - m_\tau^{\eta,c}]^Y - [m_\tau^{\eta,e} - m_\tau^{\eta,c}]^{Y-T}$	0.0198	−0.0446
$[m_\tau^{\eta,e} - m_\tau^{\eta,c}]^Y - [m_\tau^{\eta,e} - m_\tau^{\eta,c}]^{Y+bn}$	0.0692**	−0.1383*
$[m_\tau^{\eta,e} - m_\tau^{\eta,c}]^Y - [m_\tau^{\eta,e} - m_\tau^{\eta,c}]^{Y+bn-T}$	0.0761**	−0.1407*

Note: The superscripts Y, bn, T refer to labour income, benefits and taxes and national insurance, respectively. The H_0's $[m_3^{\eta,e} - m_3^{\eta,c}]^Y - [m_3^{\eta,e} - m_3^{\eta,c}]^\Gamma \leq 0$, $[m_4^{\eta,e} - m_4^{\eta,c}]^Y - [m_4^{\eta,e} - m_4^{\eta,c}]^\Gamma \geq 0$, for $\Gamma \in (\{Y-T\}, \{Y+bn\}, \{Y+bn-T\})$ are rejected at significance level 1% (***), 5% (**) or 10% (*), based on a one-tailed confidence interval procedure, implemented using a block bootstrap with 3000 replications.

moments of the distribution of shocks to the persistent component of idiosyncratic income. The effect of the policy is significant during contractions. Evaluating taxes and benefits separately, we find that, consistently, it is the benefits policy that significantly reduces risk.

Our findings regarding the beneficial impact of a social insurance policy are generally consistent with existing evidence for GB in Blundell and Etheridge (2010), for Sweden in Domeij and Floden (2010), for Germany in Busch and Ludwig (2016), for the United States in Kniesner and Ziliak (2002) and in Dynarski and Gruber (1997), and for the United States, Germany and Sweden in Busch et al. (2021), among others. Importantly, we find that for all the moments considered, the effects of benefits in reducing risk exposure are bigger than taxes and national insurance. This result coheres with evidence for the United Kingdom suggesting that benefits have stronger effects in reducing household income inequality than taxes (see figure 7a in Belfield et al. (2017) and Figures 4.5 and 4.6 in Blundell and Etheridge (2010)). In contrast, Kniesner and Ziliak (2002) find that, in the United States, the effects of taxes and transfers are quantitatively similar when studying the reduction in the variance of household earnings growth.[24]

5.1 | Effect of policy on the cyclicality of risk

In Table 6 we compare the cyclical behaviour of income risk (based on the third and fourth central moments) pre- and post-policy to assess the effectiveness of social insurance to mitigate this risk.[25] To this end, we summarise the results of tests of the hypothesis that the increase (in absolute value) in relevant moments during recessions is significantly higher post policy. In other words, we examine whether the cyclicality of risk been reduced by the social insurance policy.

The results show that taxes and national insurance contributions do not significantly reduce the cyclicality of the third and fourth central moments. In contrast, these were significantly reduced when benefits were taken into account. Busch and Ludwig (2016, 2020), using data for Germany and the United States, find that economic policy in the form of taxation and benefits policies reduces the increase in risk exposure arising from shocks to permanent income in contractions. However, they do not examine the effect of the tax and benefit policies separately.

6 | CONCLUSIONS

Using data from the BHPS (1991–2008) and a parametric econometric approach, which allowed us to exploit the history dependence of cross-sectional moments, this paper provided new evidence on the cyclical behaviour of household labour income risk in GB. We also assessed the role of social insurance policy in mitigating this source of income risk. To this end, we decomposed stochastic idiosyncratic household income into its transitory and persistent components, accounting for cohort heterogeneity over time. We focused our analysis on the cyclicality of the second to fourth moments of the probability distribution of shocks to the persistent component of income.

[24]A comparison between GB and the United States, regarding the effects of the tax and welfare system in reducing inequality, is in Blundell et al. (2018). They highlight the importance of the generosity of transfer payments in GB post-1997 in this respect.

[25]Note that because the cyclical income risk measures, based on the second moments, were generally not significantly different from zero in Table D.1, we do not test for post-policy effects in this case.

We discovered that in GB, household labour income risk increases during contractions, primarily via an increase in the absolute value of the third central moment of persistent shocks, whereas the variance remains acyclical. Taken together, the changes in the moments considered imply that households face an increased probability of receiving large negative shocks during contractions.

Households respond to increases in labour income risk via *ex-ante* precautionary and *ex-post* corrective economic activities, which lead to inequality and can affect market quantities. For example, precautionary behaviour related to higher labour income risk may lead to *ex-ante* increases in savings and labour supply as well as portfolio adjustments to include more lower-risk lower-return assets (see, e.g. Kniesner and Ziliak (2002), Low et al. (2010), and Meghir and Pistaferri (2011)). In contrast, *ex-post* responses to negative shocks to labour income might include the liquidation of assets and durable goods, changing jobs as well as family labour supply. Therefore, our results provide additional support to research that incorporates cyclicality in risk to understand economic outcomes.[26]

Risk exposure matters for household welfare and motivates government intervention to provide insurance. Cyclicality of risk, hence, suggests that increased intervention is justified during recessions. Our findings regarding the effects of economic policy imply that it has reduced both the level of risk exposure and its increase during contractions when risk exposure is measured by the magnitude of the third and fourth central moments of the distribution of persistent labour income shocks. In particular, we find that this reduction is achieved via the benefits as opposed to tax policies. This result confirms the importance of benefits as a policy instrument in mitigating income volatility, in addition to inequality, which has been previously noted by other UK studies using different methods than those employed here (see Blundell and Etheridge (2010) and Belfield et al. (2017)).

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[26]See Constantinides and Duffie (1996) and Storesletten et al. (2007) for the importance of countercyclical variance of earnings shocks for economic outcomes, as well as Mankiw (1986), Brav et al. (2002), Krebs (2007), and McKay (2017) for the significance of countercyclical left-skewness. The literature has also examined the relationship between idiosyncratic, firm-level shocks and aggregate fluctuations, as well as their joint determination (see Bloom et al. (2018)).

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SUPPORTING INFORMATION

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