CESifo WORKING PAPERS

7594
2019

April 2019

Cyclical income risk in Great Britain

Konstantinos Angelopoulos, Spyridon Lazarakis, James Malley

CESifo

Impressum:

CESifo Working Papers
ISSN 2364-1428 (electronic version)
Publisher and distributor: Munich Society for the Promotion of Economic Research - CESifo GmbH
The international platform of Ludwigs-Maximilians University's Center for Economic Studies and the ifo Institute
Poschingerstr. 5, 81679 Munich, Germany
Telephone +49 (0)89 2180-2740, Telefax +49 (0)89 2180-17845, email office@cesifo.de
Editor: Clemens Fuest
www.cesifo-group.org/wp

An electronic version of the paper may be downloaded
· from the SSRN website: www.SSRN.com
· from the RePEc website: www.RePEc.org
· from the CESifo website: www.CESifo-group.org/wp

CESifo *Working Paper No. 7594*
Category 4: Labour Markets

Cyclical income risk in Great Britain

Abstract

This paper establishes new evidence on the cyclical behaviour of household income risk in Great Britain and assesses the role of social insurance policy in mitigating against this risk. We address these issues using the British Household Panel Survey (1991-2008) by decomposing stochastic idiosyncratic income into its transitory, persistent and fixed components. We then estimate how income risk, measured by the variance and the skewness of the probability distribution of shocks to the persistent component, varies between expansions and contractions of the aggregate economy. We first find that the volatility and left-skewness of these shocks is a-cyclical and counter-cyclical respectively. The latter implies a higher probability of receiving large negative income shocks in contractions. We also find that while social insurance (tax-benefits) policy reduces the levels of both measures of risk as well as the counter-cyclicality of the asymmetry measure, the mitigation effects work mainly via benefits.

JEL-Codes: D310, E240, J310.

Keywords: household income risk, social insurance policy, aggregate fluctuations.

Konstantinos Angelopoulos
University of Glasgow
Glasgow / United Kingdom
Konstantinos.Angelopoulos@glasgow.ac.uk

Spyridon Lazarakis
University of Glasgow
Glasgow / United Kingdom
s.lazarakis.1@research.gla.ac.uk

James Malley
University of Glasgow
Glasgow / United Kingdom
Jim.Malley@glasgow.ac.uk

March 31, 2019
We would like to thank participants at the Society for Economic Dynamics 2017, Asian Meeting of the Econometric Society 2017, workshop participants at Athens University of Economics & Business and Zurich University 2018, Christopher Busch, Fabrice Collard, Thomas Crossley, Campbell Leith, Stephen Jenkins, Hamish Low, Edward Prescott, Kjetil Storesletten and Ulrich Woitek for helpful comments and suggestions on earlier versions. We are grateful for financial support from the Economic and Social Research Council, Scottish Government, as well as the Scottish and Royal Economic Societies. The usual disclaimer applies.

1 Introduction

This paper provides new evidence on the cyclical behaviour of household income risk in Great Britain (GB) from 1991-2008 and assesses 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, persistent and fixed components.[1] We measure risk by the second and the third central moment of the probability distribution of shocks to the persistent component. Following the parametric approach of Storesletten *et al.* (2004) and Busch and Ludwig (2016), we allow these two moments to depend on the aggregate state and, in particular, to vary between expansions and contractions of the aggregate economy. The advantage of this methodology is that we can identify the differences between booms and slumps by exploiting history dependent cross-sectional moments that incorporate aggregate shocks outside the panel data sample period.

Idiosyncratic risk has implications for the household, the aggregate economy and social insurance policy. Our interest in investigating the cyclical variation in third moments is motivated by recent evidence for a number of countries, suggesting that the skewness of the distribution of the growth of earnings is counter-cyclical (see e.g. Guvenen *et al.* (2014) and Busch *et al.* (2018)), and data for GB suggesting that measures of skewness of this distribution exhibit significant time variation. In particular, a preliminary investigation of the distribution of the growth of annual labour income across households, using British Household Panel Survey (BHPS) data, reveals that the third central moment has significantly more variability over time than the variance of the same distribution. We plot both series in Figure 1, where the grey shaded areas show periods that, according to the OECD measure of Composite Leading Indicators, correspond to contractions of the aggregate UK economy.[2]

[Figure 1]

First note in Figure 1 that the overall time evolution of the variance is similar with that reported in Blundell and Etheridge (2010) for shocks to the permanent component of household earnings. This variation reveals little

[1] See, e.g. Meghir and Pistaferri (2011) for a review of the earnings dynamics models.

[2] Details relating to the data, sample selection, and variable definitions used in Figures 1-2 and throughout the paper are reported in Appendix A. Note that we use the third central moment in Figures 1-2 to correspond to what we estimate in Tables 1-6 below. Also note that the conclusions we draw from these Figures and Tables are qualitatively the same when we instead calculate the standardised third moments (see Appendix C). In light of this, we use the term skewness when referring to both the third central moment and the standardised third moment or Pearson's moment coefficient of skewness.

relationship with the aggregate state, which is also broadly consistent with previous results for wage risk in GB in Bayer and Juessen (2012). The time variation of the third central moment, on the other hand, is suggestive of a more volatile distribution of income growth with respect to its asymmetry. In addition, we note that, at least after 1996, the general pattern of the third central moment indicates rises in periods of expansions and falls in periods of contraction. However, the relationship with the aggregate state is not robust. Moreover, since it refers to changes to the overall labour income, we cannot infer from this whether persistent shocks, which have stronger effects on the household (see e.g. Meghir and Pistaferri (2011) and references therein), exhibit cyclicality. The relative importance of these shocks motivates our interest in whether cyclical variation is present in the third moments of shocks to the permanent component of household income. Busch and Ludwig (2016) find evidence of such cyclicality in Germany.

An evaluation of the extent to which different components of economic policy reduce risk exposure, and in particular the increase in vulnerability during contractions, is important in informing policy making, as it suggests which policy instrument is likely to be more effective when insurance is most required. The evidence for Germany in Busch and Ludwig (2016) and for US, Germany and Sweden in Busch *et al.* (2018) shows that social insurance policy does reduce the increase in risk exposure associated with changes in skewness. However, it also suggests that, at least when looking at distributions of growth in earnings (Busch *et al.* (2018)), this is mainly driven by taxes, compared with transfers. On the other hand, evidence from the UK (see e.g. Belfield *et al.* (2017)) suggests that benefits have a stronger effect in mitigating an increase in inequality, especially in contractions. Moreover, when we plot the variance and third moment of different measures of pre- and post-government household income growth in GB, in Figure 2, we also see that benefits seem to be dominant in smoothing the time variation of household income risk.

[Figure 2]

Figure 2 shows that policy does reduce the level of second and third moments in the first and second subplots respectively, as well as their volatility over time. Moreover, these effects are driven primarily by benefits. This evidence further motivates our interest in whether these effects maintain when evaluating the ability of policy to provide insurance by smoothing the cyclicality of shocks to permanent risk.

To assess whether cyclical variation is present in the second and third moments of shocks to the permanent component of household income and whether these effects remain post policy we use the British Household Panel

Survey (BHPS). This dataset has been used extensively for income dynamics analysis in the UK (see, e.g. Blundell and Etheridge (2010), Bayer and Juessen (2012), Capellari and Jenkins (2014) and Etheridge (2015)), since it provides measures of annual earnings at the individual and household levels, in addition to observable characteristics. The latter allow us to partial out observable deterministic components (i.e. time, experience, education, region of residence and household size effects) to isolate idiosyncratic labour income in the data. We then employ an estimation procedure that aims to match the theoretical moments of the model of state-dependent income dynamics, with their empirical counterparts. To examine the effect of social insurance, we estimate risk for different measures of pre- and post-policy household income, using the data in Bardasi *et al.* (2012) and evaluate the effect of tax and benefits policies on the level and cyclicality of risk.

We find that the volatility and left-skewness of the shocks to the permanent component of income are a-cyclical and counter-cyclical respectively. The latter implies a higher probability of receiving large negative income shocks in contractions. In addition, we find that while taxes and benefits help to reduce the levels of both measures of risk, it is benefits that significantly reduce the increase in risk exposure associated with skewness during bad times. More generally, benefits have stronger risk mitigation effects.

The rest of the paper is organised as follows. In the next section we review the relevant literature to provide the context motivating the approach used in our analysis. In Section 3 we discuss the methods employed to obtain the two measures of income risk and in Section 4 the data and econometric estimation procedure. The results are presented in Section 5 and Section 6 contains the conclusions.[3]

2 Related literature

Idiosyncratic risk matters for individual (or household) level behaviour and outcomes. In response to labour income risk, individuals or households engage in a number of *ex ante* precautionary and *ex post* corrective economic activities, which ultimately can affect aggregate economic outcomes (see e.g. Low *et al.* (2010) and Meghir and Pistaferri (2011)). For example, precautionary behaviour related to higher labour income risk may lead to increases in savings and labour supply as well as portfolio adjustments to include more lower-risk lower-return assets. These responses are stronger under incomplete markets. In contrast, *ex post* responses to negative shocks to labour

[3]Further details on the data are reported in Appendix A. Additional empirical results can also be found in Appendices C and D relating to the robustness of our findings.

income might include the liquidation of assets and durable goods, changing jobs and family labour supply, and adjustments in consumption. The effects of idiosyncratic risk are typically stronger for persistent, relative to purely transitory, income shocks. This has motivated a decomposition of income risk into persistent (or permanent) and transitory components, emphasising the importance of the former. The absence of market opportunities for insurance against negative shocks to labour income typically motivates public insurance.[4]

The cyclical behaviour of labour income risk, and in particular the extent to which risk exposure increases during periods of contraction, is thus important for individual behaviour and outcomes. Moreover, the relationship between income risk and aggregate economic conditions is also important for understanding macroeconomic phenomena. Theoretical work has focused on the role of counter-cyclical risk in explaining asset prices and economic fluctuations (see e.g. the research reviewed in Storesletten *et al.* (2004) and Guvenen *et al.* (2014)). The main idea is that idiosyncratic labour income risk is increasing with respect to negative aggregate shocks. In this literature, some studies have concentrated on the importance of the counter-cyclical variance of earnings shocks (e.g. Constantinides and Duffie (1996) and Storesletten *et al.* (2007)) while others have highlighted the significance of the counter-cyclical left-skewness of earnings shocks (e.g. Mankiw (1986), Brav *et al.* (2002), Krebs (2007), and McKay (2017)). From a policy perspective, understanding the cyclical properties of household income risk, and the extent to which social insurance can mitigate increases in risk exposure, is important for the evaluation of alternative policies.

These considerations have motivated empirical research which examines the relationship between higher moments of the distribution of individual and household labour income shocks and changes in aggregate outcomes. Given the importance of persistent income shocks, a small number of studies have directly examined the cyclical properties of the shocks to the persistent component of individual or household income. In a seminal contribution, 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 (earnings plus benefits) is counter-cyclical. Following the same approach and dataset, Bayer and Juessen (2012) find that household wages have countercyclical variance in the US, but that the

[4] Such negative shocks can take the form of unemployment or health shocks that reduce employment, or shocks that reduce returns to work, e.g. shocks that lower productivity, technology shocks that make skills less valuable and shocks leading to employer-worker mismatch.

variance of idiosyncratic shocks to wages is a-cyclical in GB, using BHPS data.[5] Busch and Ludwig (2016), using data for Germany for individuals and households, extend the approach in Storesletten *et al.* (2004) and estimate a model for income dynamics that allows for regime-switching variance and skewness. They find that both the variance and left-skewness of shocks to the permanent components of income are counter-cyclical.

A relatively larger set of studies has studied the cyclicality of the distribution of income shocks across individuals or households, approximating shocks with growth rates of relevant measures of income, without statistically decomposing shocks to those affecting the persistent and the transitory component of income. For example, Ziliak *et al.* (2011) used the US matched Current Population Survey and found that the volatility of individual male and female earnings growth are counter-cyclical and pro-cyclical respectively, whereas Cappellari and Jenkins (2014), using BHPS data, find that the variance of individual earnings growth shows little time variation over the 1991-2008 period. In an influential contribution, Guvenen *et al.* (2014), using US Social Security Administration data without imposing restrictions on the shape of the distribution of shocks to individual earnings, discover that the left-skewness is counter-cyclical and variance is a-cyclical. The counter-cyclical property of the skewness of income risk has been further documented in the literature using panel data surveys for Germany, Sweden and the US in Busch *et al.* (2018). These findings have important implications for the cyclical properties of risk exposure, as they suggest that in periods of contraction the probability of receiving large negative shocks increases. Given the importance of shocks to the persistent component of income relative to the transitory part, these results further motivate research into examining the cyclical variation of such income shocks.[6]

The literature has also examined the risk mitigation performance of social insurance policies. Regarding the cyclicality in risk exposure associated with changes in skewness of the distribution of earnings growth, Busch *et al.* (2018) provide evidence from the US, Germany and Sweden that social insurance policy does reduce the increase in risk exposure. Their results suggest that taxes have the biggest effect in reducing the cyclicality of skewness. Their finding for the US is in line with the results in Kniesner and Zilliak (2002) who show that taxes have similar effects in reducing the variance of

[5] BHPS data have been frequently used to decompose earnings risk into its transitory and persistent components in Great Britain (see also Blundell and Etheridge (2010) and Etheridge (2015)).

[6] Guvenen *et al.* (2014) approximated permanent shocks by 5-year differences in income. However, for shorter time series, this approach becomes more difficult and a statistical decomposition as in Busch and Ludwig (2016) is required.

the distribution of household earnings growth than transfers. The effects of the tax-benefit social insurance system in reducing the variance of permanent and transitory income shocks has been demonstrated in e.g. Blundell and Etheridge (2010) for GB and in Domeij and Floden (2010) for Sweden. However, evidence from e.g. Blundell and Etheridge (2010) and Belfield *et al.* (2017) demonstrates that, in the UK benefits have stronger effects than taxes in mitigating household income inequality.

Regarding risk mitigation associated with the cyclicality of the third moment of shocks to the persistent component of income, Busch and Ludwig (2016) show that the tax and transfer system in Germany reduces the increase in risk exposure arising from shocks to permanent income in contractions. However, they do not disaggregate the effects of taxes and transfers. The importance of shocks to the persistent component of income relative to the transitory part motivates us to further investigate which policy instrument is likely to be more effective in mitigating the increase in risk exposure in periods of contraction.

3 Idiosyncratic Income Risk

We next set out the method used to derive the measures of idiosyncratic income risk which vary depending on the aggregate state of the economy. 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 time period 1991-2008.

3.1 Idiosyncratic income shocks

Following the literature on modeling earnings dynamics (see e.g. Meghir and Pistaferri (2011)), we assume that idiosyncratic component of income, $\mu_{i,h,t}$, for household i of age h, $h \in \{1, 2, ..., H = 36\}$, in period t, $t \in \{1, 2, ..., T = 18\}$, is driven by stochastic fixed effects, χ_i, persistent effects, $z_{i,h,t}$, and transitory shocks, $\varepsilon_{i,t}$:

$$\mu_{i,h,t} = \chi_i + z_{i,h,t} + \varepsilon_{i,t}, \tag{1}$$

$$z_{i,h,t} = \rho z_{i,h-1,t-1} + \eta_{i,t}, \tag{2}$$

where $0 < \rho < 1$ and $\eta_{i,t}$ captures innovations to the persistent effects.

Following Busch and Ludwig (2016), the distributional assumptions for the three components in (1) and (2) are:

$$\chi_i \underset{i.i.d.}{\sim} F_\chi(0, m_2^\chi, m_3^\chi), \tag{3}$$
$$\varepsilon_{i,t} \underset{i.i.d.}{\sim} F_\varepsilon(0, m_2^\varepsilon, m_3^\varepsilon), \tag{4}$$
$$\eta_{i,t} \underset{i.i.d.}{\sim} F_\eta\left(0, m_2^{\eta,f(t)}, m_3^{\eta,f(t)}\right), \tag{5}$$

where F_χ, F_ε, and F_η denote the density functions of χ_i, $\varepsilon_{i,t}$ and $\eta_{i,t}$ respectively. All the moments for the fixed effects, (m_2^χ, m_3^χ) and the transitory shocks, $(m_2^\varepsilon, m_3^\varepsilon)$, are constant. In contrast, the innovations to the persistent shocks, $\left(m_2^{\eta,f(t)}, m_3^{\eta,f(t)}\right)$ are assumed to be time dependent since we wish to test whether they are driven by the aggregate state of the economy.

Furthermore, following e.g. Storesletten *et al.* (2004), we assume that it is the history of persistent shocks only after the age $h = 1$ that matters for idiosyncratic income. In particular, we assume that $z_{i,0,t} = 0$, implying that prior to joining the labour market there are no persistent shocks that matter for earnings dynamics after $h = 1$ other than the fixed effects. In other words, the fixed effects capture factors that matter for income dynamics prior to joining the labour market.

Following Storesletten *et al.* (2004) and Busch and Ludwig (2016) we allow $m_2^{\eta,f(t)}$ and $m_3^{\eta,f(t)}$ to take two values each depending on the aggregate state, $f(t)$, which is either an expansion, e, or a contraction, c. For example, we define an indicator variable $I_{f(t)=e}$ to be equal to 1 if period t is an expansion and $I_{f(t)=c}$ to be equal to 0 if period t is a contraction, i.e.:

$$m_2^{\eta,f(t)} \equiv \left(I_{f(t)=e}\right) m_2^{\eta,e} + \left(1 - I_{f(t)=c}\right) m_2^{\eta,c}, \tag{6}$$
$$m_3^{\eta,f(t)} \equiv \left(I_{f(t)=e}\right) m_3^{\eta,e} + \left(1 - I_{f(t)=c}\right) m_3^{\eta,c}. \tag{7}$$

3.2 Theoretical moments

The above assumptions imply a particular structure on the covariance matrix of the stochastic processes $z_{i,h,t}$ and $\mu_{i,h,t}$. Applying expected value, variance, co-variance and co-skewness rules to the above model for $\mu_{i,h,t}$ yields the following set of theoretical moments which will be employed in our empirical analysis:[7]

$$E\left(\mu_{i,h,t}\right) = 0, \tag{8}$$
$$E\left(\mu_{i,h,t}^2\right) = m_2^\chi + m_2^\varepsilon + E\left(z_{i,h,t}^2\right), \tag{9}$$

[7] We use the convention here that $h = 1$ when the age is 25, and goes through to $h = 36$ when the age is 60.

$$E\left(\mu_{i,h,t}^3\right) = m_3^\chi + m_3^\varepsilon + E\left(z_{i,h,t}^3\right), \tag{10}$$

$$\begin{aligned} Cov(\mu_{i,h,t}, \mu_{i,h+\kappa,t+\kappa}) &= E(\mu_{i,h,t}\mu_{i,h+\kappa,t+\kappa}) \\ &= m_2^\chi + E\left(z_{i,h,t}^2 z_{i,h+\kappa,t+\kappa}\right), \end{aligned} \tag{11}$$

$$\begin{aligned} CoSk(\mu_{i,h,t}, \mu_{i,h+\kappa,t+\kappa}) &= E(\mu_{i,h,t}^2, \mu_{i,h+\kappa,t+\kappa}) \\ &= m_3^\chi + E\left(z_{i,h,t}^2 z_{i,h+\kappa,t+\kappa}\right), \end{aligned} \tag{12}$$

where

$$\begin{aligned} E\left(z_{i,h,t}^2\right) &= \sum_{j=0}^{h-1} \rho^{2j} m_2^{\eta,(f(t-j))}, \\ E\left(z_{i,h,t}^3\right) &= \sum_{j=0}^{h-1} \rho^{3j} m_3^{\eta,(f(t-j))}, \\ E\left(z_{i,h,t} z_{i,h+\kappa,t+\kappa}\right) &= \rho^\kappa \sum_{j=0}^{h-1} \rho^{2j} m_2^{\eta,(f(t-j))}, \\ E\left(z_{i,h,t}^2 z_{i,h+\kappa,t+\kappa}\right) &= \rho^\kappa \sum_{j=0}^{h-1} \rho^{3j} m_3^{\eta,(f(t-j))}. \end{aligned}$$

The moments of the idiosyncratic component of income, $\mu_{i,h,t}$, are thus a function of past moments of innovations to the persistent component. Therefore, the estimation of the parameters of interest requires knowledge of whether $H-1$ years prior to those in the observed sample of the 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).

3.3 Empirical moments

To obtain the idiosyncratic component of household income, $\mu_{i,h,t}$, we follow the literature on earnings dynamics and run a Mincerian-type regression to partial out non-stochastic effects from labour income. In particular, we assume that the process determining the logarithm of annual household income, $y_{i,h,t}$, is comprised of an observable deterministic part, $d_t + bx_{i,h,t}$, and the unobservable random component, $\mu_{i,h,t}$:

$$y_{i,h,t} = d_t + bx_{i,h,t} + \mu_{i,h,t}, \tag{13}$$

where b is a vector of parameters. In particular, the regressors in (13) include calendar year time effects, d_t, and a set of dummy variables, $x_{i,h,t}$, for experience (approximated by age), region of residence 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 (13) for two separate skill groups, i.e. households whose head has University education and those households whose head does not. Finally, since in our econometric analysis we employ household quantities for the arguments in (13), we define the age and regional effects in terms of the head of the household. We denote by $\widehat{\mu}_{i,h,t}$ the estimated idiosyncratic component of household income.

Using a panel dataset of household incomes for time $t \in \{1, 2, ..., T\}$, age $h \in \{1, 2, ..., H\}$ and $i \in \{1, 2, ..., N\}$ we first calculate the empirical moments of the idiosyncratic income shocks using the residuals from the Mincer regression. In particular, every year t, we group agents in the sample into 5-year adjacent age cells indexed by h, i.e. we define an individual or a household as belonging to the age group h if her true age was between $h - 2$ and $h + 2$. For example, the first cell, i.e. age group 25, contains all workers between 23 and 27 years old, the second cell, i.e. age group 26, contains all workers between 24 and 28 years old, while the last cell, i.e. age group 60, contains all workers between 58 and 62 years old. Our sample length and age grouping imply $T = 18$ and $H = 36$ which implies a total of $2 \times 5,187$ empirical moments. In particular, the empirical moments are given by:

$$\begin{aligned} &\tfrac{1}{I_{h,t,\kappa}} \textstyle\sum_{i=1}^{N} \iota_{i,h,t,\kappa} \left[\left(\widehat{\mu}_{i,h,t}\right)^{\phi} \left(\widehat{\mu}_{i,h+\kappa,t+\kappa}\right)^{\psi} \right] = \\ &= \tfrac{1}{I_{h,t,\kappa}} \textstyle\sum_{i=1}^{N} \iota_{i,h,t,\kappa} \Big[\left(y_{i,h,t} - \widehat{d}_t - \widehat{b} x_{i,h,t} \right)^{\phi} \times \\ &\times \left(y_{i,h+\kappa,t+\kappa} - \widehat{d}_t - \widehat{b} x_{i,h+\kappa,t+\kappa} \right)^{\psi} \Big], \end{aligned} \tag{14}$$

where $(\phi, \psi) \in \{(1,1), (2,1)\}$, $\kappa = 0, ..., \min[T - t, H - h]$, $I_{h,t,\kappa} = \sum_{i=1}^{N} \iota_{i,h,t,\kappa}$ 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.

4 Data and estimation

In this section we provide information on the dataset and variables used for the analysis as well as a brief description of the sample selection criteria, followed by a description of the econometric methods used to estimate the model parameters.

4.1 Data

The BHPS is a comprehensive longitudinal study for GB, covering 1991 to 2008.[8] It 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, as well as on socio-economic characteristics of the respondents. These characteristics include gender, education, age, social (professional) class and region.[9] BHPS was replaced in 2010 by a new panel data survey, Understanding Society, which however does not include information on annual earnings, and thus cannot be used to analyse earnings risk. We also make use of 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 Bardasi *et al.* (2012) dataset tracks the same individuals/households for the same time as the BHPS i.e. 1991-2008.

4.1.1 Household level

We start with the allocation of individuals to households from BHPS and keep households with a spouse/partner relationship (hence discarding those households comprised of a single member or those that involve cohabiting but not family-related members) as well as those where the head is between 23-62 years and reports non-zero labour income.[10] Following e.g. Blundell and Etheridge (2010) we define the head to be the older married (or in partnership) male. We also have measures on annual earnings of the household's individual members.

Using the DCANHIV dataset we have consistent series of household labour income, gross income, gross income less taxes and national insurance contributions, gross income plus benefits, and gross income plus benefits less taxes and national insurance contributions. Labour income is the sum of annual earnings of the household members. Gross income is equal to household's

[8]Further details on the datasets and the definition and construction of variables and information on sample selection can be found in Appendix A.

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

[10]Some households defined as such have additional members, e.g. other family members living in the same household.

labour income plus annual investment income, occupational pension income and annual private transfers income. Taxes are the annual household income taxes after credits, while benefits are the annual social benefits income, which totals all receipts from state benefits from all household's members (including national insurance retirement pensions).

To ensure strong attachment to the labour marker, we follow e.g. Guvenen *et al.* (2014), Busch *et al.* (2016) and include in any year households in which their head reports annual earnings greater than half of the product between the minimum legal hourly wage times 520 hours, implying at least a few months of work during the year. For each year, we order the households according to their labour income and we discard the observations who are in the top 1%.

4.1.2 Aggregate Shocks

As a proxy for the aggregate state of the economy, we use the OECD Composite Leading Indicators (CLI) for the United Kingdom "from the peak through the trough" which can be found in Fred St. Louis website.[11] The OECD identifies months of turning points without designating a date within the month that the turning points occurred. The dummy variable adopts an arbitrary convention that the turning point occurred at a specific date within the month. To be consistent with the BHPS data, we have chosen the annual frequency and as an aggregation period the end of period (from September to September). We aggregate on the monthly indices and set as contractions the years with 6 or more months of contraction.[12] The OECD based aggregate cycle indicator can be extended into the past until 1956 which corresponds with the year (i.e. 1991) that the oldest individuals in the sample entered the labour market at age 25.

4.2 Estimation

The moment conditions employed in the GMM estimation are:

[11] The components of the CLI are time series which exhibit leading relationships with the reference series (GDP) at turning points. Country CLIs are compiled by combining de-trended smoothed and normalized components. The component series for each country are selected based on various criteria such as economic significance; cyclical behaviour; data quality; timeliness and availability.

[12] Note that, alternatively, following Busch and Ludwig's (2016) method to characterise years as either contractionary or expansionary periods gives us exactly the same classification.

$$E\left[\widehat{\mu}^2_{i,h,t} - \mu^2_{i,h,t}(\theta)\right] = 0, \quad (15)$$
$$E\left[\widehat{\mu}^3_{i,h,t} - \mu^3_{i,h,t}(\theta)\right] = 0, \quad (16)$$
$$E\left[\widehat{\mu}_{i,h,t}\widehat{\mu}_{i,h+\kappa,t+\kappa} - \mu_{i,h,t}(\theta)\,\mu_{i,h+\kappa,t+\kappa}(\theta)\right] = 0, \quad (17)$$
$$E\left[\widehat{\mu}^2_{i,h,t}\widehat{\mu}_{i,h+\kappa,t+\kappa} - \mu^2_{i,h,t}(\theta)\,\mu_{i,h+\kappa,t+\kappa}(\theta)\right] = 0, \quad (18)$$

where θ is the vector of parameters to be estimated:

$$\theta = \{\rho, m_2^\chi, m_3^\chi, m_2^\varepsilon, m_3^\varepsilon, m_2^{\eta,c}, m_2^{\eta,e}, m_3^{\eta,c}, m_3^{\eta,e}\}.$$

The empirical moments in conjunction with the theoretical ones given by (8)-(12) allow us to identify: (i) the persistence parameter ρ; (ii) the second and third moments of distribution of the fixed effects, m_2^χ and m_3^χ; (iii) the second and third moments of distribution of the transitory shocks, m_2^ε and m_3^ε; and (iv) the time dependent higher moments for innovations to the persistent component i.e. $m_2^{\eta,f(t)}$, $m_3^{\eta,f(t)}$. We show analytically in Appendix B that the parameters in θ can be identified if we have at least four time periods and four age groups.[13] In the data, we have 18 periods and 36 age groups, which implies that the system (15)-(18) is over-identified.

Let $\mathbf{m}$ be the vector with all the available empirical moments constructed as above and $\mathbf{G}(\theta)$ the vector of the respective theoretical moments. The goal is to estimate a model for **m:**

$$\mathbf{m} = \mathbf{G}(\theta) + \boldsymbol{\Upsilon}, \quad (19)$$

where $\boldsymbol{\Upsilon}$ captures sampling variability. For the estimation, we minimize the distance between the empirical and the theoretical moments. Formally, we numerically minimize the following objective function:

$$Q(\theta) = \min_\theta \left(\mathbf{m} - \mathbf{G}(\theta)\right)' \mathcal{W} \left(\mathbf{m} - \mathbf{G}(\theta)\right), \quad (20)$$

where $\mathcal{W}$ is a weighting matrix. Following Altonji and Segal (1996), the typical choice of $\mathcal{W}$ in the literature is the identity matrix. However, notice that each moment is calculated by a different number of observations. Moreover,

[13] In particular, in Appendix B, we illustrate how to identify the persistence parameter, ρ, using the minimum number of consecutive time periods and age groups, i.e. 4 for each. Using equation, (14), this example implies 60 empirical moments. The Appendix also illustrates, conditional on a given value of ρ, how to identify: m_2^χ, m_3^χ, m_2^ε, m_3^ε, $m_2^{\eta,f(t)}$ and $m_3^{\eta,f(t)}$, again using the minimum number of consecutive time periods and age groups, i.e. 3 for each. Using equation, (14), this example implies 28 empirical moments.

since we are calculating higher moments, it is well known that bigger samples give more accurate results. Hence, we weight each moment equation by the number of observations used to calculate its empirical part since the panel is unbalanced.[14]

To compute the standard errors, we follow MaCurdy (2007), and use the block bootstrap procedure for 1000 replications. The resulting confidence intervals account for serial correlation of arbitrary form, heteroskedasticity as well as for the fact that we use pre-estimated residuals.[15] Formally, the bootstrap p-values for an estimator θ are calculated as:

$$2 * \left[1 - \Phi\left(\frac{\widehat{\theta}}{\sigma_{\widehat{\theta}}}\right)\right], \tag{21}$$

where $\widehat{\theta}$ is the GMM estimator and $\sigma_{\widehat{\theta}}$ its bootstrap standard errors. Φ denotes the Normal cumulative distribution function.

5 Results

The estimated parameters $\theta = [\rho, m_2^{\chi}, m_3^{\chi}, m_2^{\varepsilon}, m_3^{\varepsilon}, m_2^{\eta,c}, m_2^{\eta,e}, m_3^{\eta,c}, m_3^{\eta,e}]$ are reported in Tables 1 and 2. Table 1 concentrates on the main parameters of interest, i.e. the second and third moments of the probability distribution of shocks to the persistent component of idiosyncratic income during expansions and contractions. Whereas, Table 2 reports the remaining parameter estimates relating to the transitory shocks and stochastic fixed effects. Tables 3-6 then report the results of statistical tests relating to the effect of tax and benefit policy on the levels of persistent and transitory income risk as well as on the cyclicality of persistent income risk.

5.1 Cyclical risk

Columns 1-2 and 4-5 in Table 1 present estimates of $[m_2^{\eta,c}, m_2^{\eta,e}, m_3^{\eta,c}, m_3^{\eta,e}]$ for households across five different measures of labour income. Columns 3 and 6 in Table 1 also report the difference between each moment in expansions and in contractions to test whether income risk increases in bad times.

[14] For similar treatment see Heathcote *et al.* (2010) and Domeij and Floden (2010).
[15] See also Hall and Horowitz (1996) and Horowitz (2003).

Table 1: Cyclical household income risk

	$m_2^{\eta,e}$	$m_2^{\eta,c}$	$m_2^{\eta,e} - m_2^{\eta,c}$	$m_3^{\eta,e}$	$m_3^{\eta,c}$	$m_3^{\eta,e} - m_3^{\eta,c}$
			labour income			
est	0.0370***	0.0389***	-0.0019	-0.0105	-0.0336***	0.0231*
s.e.	(0.0110)	(0.0128)	(0.0073)	(0.0076)	(0.0124)	(0.0154)
			gross income			
est	0.0346***	0.0402***	-0.0056	-0.0049	-0.0373***	0.0324***
s.e.	(0.0083)	(0.0095)	(0.0072)	(0.0071)	(0.0098)	(0.0126)
			gross income −taxes −NI			
est	0.0286***	0.0329***	-0.0043	-0.0031	-0.0273***	0.0242***
s.e.	(0.0065)	(0.0079)	(0.0061)	(0.0051)	(0.0076)	(0.0099)
			gross income +benefits			
est	0.0240***	0.0267***	-0.0027	-0.0005	-0.0076**	0.0072*
s.e.	(0.0046)	(0.0068)	(0.0050)	(0.0031)	(0.0032)	(0.0050)
			gross income +benefits −taxes −NI			
est	0.0185***	0.0214***	-0.0029	0.0001	-0.0049**	0.0051*
s.e.	(0.0038)	(0.0060)	(0.0042)	(0.0022)	(0.0023)	(0.0036)

The Bootstrap standard errors are included in parentheses (1000 sims) and the Bootstrap p-values are denoted as $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Note that the p-values for all columns are for a two-tailed test, except the difference tests which are for a one-tailed test. The notation adopted here for the statistical tests apply throughout the paper. NI refers to national insurance contributions.

5.2 Second moments

Columns 1 and 2 in Table 1 reveal statistically significant second moments, m_2^{η}, across all five income measures in both expansions and contractions. Counter-cyclical volatility, implies that income risk is higher in contractions than in expansions, i.e. $m_2^{\eta,e} < m_2^{\eta,c}$ or that $m_2^{\eta,e} - m_2^{\eta,c} < 0$. The signs associated with the difference, $m_2^{\eta,e} - m_2^{\eta,c}$, in column 3 of Table 1 qualitatively suggest counter-cyclical volatility for all measures. However, a one-sided test of the null hypothesis that $H_0 : m_2^{\eta,e} - m_2^{\eta,c} \geq 0$ against the alternative hypothesis that $H_A : m_2^{\eta,e} - m_2^{\eta,c} < 0$ implies an a-cyclical volatility for all measures since the null cannot be rejected for any case considered.

These results broadly cohere with the finding of Bayer and Juessen (2012) who also find a-cyclical volatility of wage risk for the UK. However, it should be noted that our results are not directly comparable to Bayer and Juessen (2012) given that we employ different measures of household compensation. In particular, we use labour income which includes both wages and employment for households, whereas Bayer and Juessen (2012) use the average hourly wage of the head and spouse. Nonetheless, evidence from both papers points in the same direction. 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 shocks components over 1991-2003 in their Figure 6.1 do not show evident co-movement with the aggregate conditions.

5.3 Third moments

Column 4 in Table 1 shows statistically insignificant third moments, m_3^η, across all five income measures in expansions. Whereas, column 5 in Table 1 shows statistically significant (negatively signed) third moments, m_3^η, for all five income measures in contractions. Several observations regarding the third moment results are worth pointing out. First, the time variation in m_3^η between expansions and contractions implies that the idiosyncratic income shocks are clearly drawn from a non-normal distribution. Second, two distributions of income shocks with the same variance can imply very different amounts of risk if they differ in m_3^η. For example, the asymmetry in the distribution of idiosyncratic income shocks implied by a non-zero third moment suggests that, depending on its sign, one of the two tails of the distribution is longer.

A negative third moment signifies that the distribution is skewed to the left and the left tail is longer than the right tail. In our case, since the left tail represents the bad shocks to income, a longer left tail in contractions than in expansions implies that there is a higher probability of a household receiving a large negative income shock in bad times. Thus, income risk which is higher in contractions than in expansions, i.e. $m_3^{\eta,e} > m_3^{\eta,c}$ or that $m_3^{\eta,e} - m_3^{\eta,c} > 0$ can be characterised as pro-cyclical asymmetry. Note that Guvenen *et al.* (2014) refers to this relationship as counter-cyclical left-skewness since left-skewness is simply defined as the negative of skewness. Nonetheless, the interpretation is the same, in a contraction the third moment is smaller (i.e. more negative) than in an expansion.

A one-sided test of the null hypothesis that $H_0 : m_3^{\eta,e} - m_3^{\eta,c} \leq 0$ against the alternative hypothesis that $H_A : m_3^{\eta,e} - m_3^{\eta,c} > 0$ suggests a significant pro-cyclical asymmetry or counter-cyclical left-skewness across all five income measures since we can reject the null in all cases considered. This constitutes new evidence for GB and coheres with international evidence, as discussed in Section 2. Notably, Busch and Ludwig (2016) is the only other study that we are aware of that explicitly decomposes shocks to earnings to investigate statistically the counter-cyclicality of third moments of shocks to the permanent component separately from possible fixed effects and transitory shocks. Although we allow for persistent, as opposed to permanent shocks (see Table 2 below for evidence in support of this), our findings for GB are similar to those in Busch and Ludwig (2016) for Germany.

Table 2: Persistence and remaining moments

	ρ	m_2^{ε}	m_3^{ε}	m_2^{χ}	m_3^{χ}
			labour income		
est	0.8530***	0.0320**	-0.0441***	0.0556***	-0.0156
s.e.	(0.0580)	(0.0132)	(0.0066)	(0.0136)	(0.0122)
			gross income		
est	0.8358***	0.0268***	-0.0339***	0.0654***	-0.0152*
s.e.	(0.0493)	(0.0081)	(0.0049)	(0.0113)	(0.0090)
			gross income −taxes −NI		
est	0.8473***	0.0239***	-0.0233***	0.0521***	-0.0099
s.e.	(0.0450)	(0.0068)	(0.0037)	(0.0096)	(0.0071)
			gross income + benefits		
est	0.8572***	0.0206***	-0.0116***	0.0487***	-0.0059
s.e.	(0.0394)	(0.0059)	(0.0021)	(0.0083)	(0.0041)
			gross income +benefits −taxes −NI		
est	0.8697***	0.0188***	-0.0064***	0.0361***	-0.0031
s.e.	(0.0397)	(0.0056)	(0.0017)	(0.0067)	(0.0033)

Note that the p-values for all columns are for a two-tailed test.

5.4 Persistence and remaining moments

Table 2 above presents the results relating to the AR(1) parameter, ρ, for the persistent component of idiosyncratic income given in equation (2) as well as the second and third moments of the probability distribution of shocks to the transitory and fixed effects parts of idiosyncratic income in equations (3) and (4) respectively. These result suggest that the vast preponderance of the 30 parameter estimates are significantly different from zero at the 1% level of significance. Only the third moment of shocks to stochastic fixed effects appears to play little role in the estimation.

The estimates reveal that the distribution of transitory shocks is also skewed to the left and the left tail is longer than the right tail. Finally, the estimate for the persistence parameter is significantly lower than 1, implying that shocks to the persistent component have high persistence but are not permanent.

5.5 Effects of policy on the levels of risk

Using row-wise comparisons of the moment estimates reported in Table 1, starting with gross income, we next test whether policy significantly lowers the levels of risk that households face. To this end, Table 3 reports the results of a one-sided test of the null hypothesis that $H_0 : [m_2^{\eta,f(t)}]^g \leq [m_2^{\eta,f(t)}]^{pp}$ against the alternative hypothesis that $H_A : [m_2^{\eta,f(t)}]^g > [m_2^{\eta,f(t)}]^{pp}$, where $f(t) = e$ for expansions and $f(t) = c$ for contractions; g refers to gross income; and *pp* refers to post-policy income. Table 4 repeats this test for the third moments.

Table 3: Policy effects on income risk (volatility)

	expansions	contractions
[1]	$[m_2^{\eta,e}]^g - [m_2^{\eta,e}]^{g-t-ni}$	$[m_2^{\eta,c}]^g - [m_2^{\eta,c}]^{g-t-ni}$
est	0.0060	0.0073
s.e.	(0.0105)	(0.0124)
[2]	$[m_2^{\eta,e}]^g - [m_2^{\eta,e}]^{g+b}$	$[m_2^{\eta,c}]^g - [m_2^{\eta,c}]^{g+b}$
est	0.0106	0.0135
s.e.	(0.0095)	(0.0117)
[3]	$[m_2^{\eta,e}]^g - [m_2^{\eta,e}]^{g+b-t-ni}$	$[m_2^{\eta,c}]^g - [m_2^{\eta,c}]^{g+b-t-ni}$
est	0.0161**	0.0189**
s.e.	(0.0091)	(0.0113)

The superscripts g, b, t and ni refer to gross income, benefits, taxes, and national insurance respectively in Tables 3-6.

The positive differences between gross income and the various measures of income net of policy reported in Table 3 indicate, qualitatively, that tax and benefit policy is working in the right direction and reduces the spread of the distribution of shocks to the persistent component of idiosyncratic income. However, when considering the effects of public insurance, only gross income plus benefits net of taxes and national insurance is statistically significant in both expansions and contractions (see row [3]). In other words, it is the combination of taxes and benefits that reduces the variance of risk, in either aggregate state. The results regarding the overall effect of social insurance policy in reducing the level of the variance of shocks to income are consistent with the findings in Blundell and Etheridge (2010), who also find big reductions in the variance of shocks to the permanent component of household income when comparing household earnings with disposable income (see their figure 6.1).

Turning to Table 4 we can see that public policy has not statistically significantly reduced the level of income risk reflected by left-skewness during expansions. This is not surprising since we learned from the fourth column in Table 1 that during expansions the third central moments are not significantly different from zero. However, during contractions, public benefits on their own and public benefits net of taxes and national insurance have significantly reduced this level of risk (see rows [2] and [3] respectively in Table 4).

Table 4: Policy effects on income risk (asymmetry)

	expansions	contractions
[1]	$[m_3^{\eta,e}]^g - [m_3^{\eta,e}]^{g-t-ni}$	$[m_3^{\eta,c}]^g - [m_3^{\eta,c}]^{g-t-ni}$
est	-0.0018	-0.0100
s.e.	(0.0087)	(0.0125)
[2]	$[m_3^{\eta,e}]^g - [m_3^{\eta,e}]^{g+b}$	$[m_3^{\eta,c}]^g - [m_3^{\eta,c}]^{g+b}$
est	-0.0044	-0.0296***
s.e.	(0.0078)	(0.0102)
[3]	$[m_3^{\eta,e}]^g - [m_3^{\eta,e}]^{g+b-t-ni}$	$[m_3^{\eta,c}]^g - [m_3^{\eta,c}]^{g+b-t-ni}$
est	-0.0050	-0.0323***
s.e.	(0.0076)	(0.0101)

Finally, using a row-wise comparisons of the moment estimates reported in Table 2, starting with gross income, we next test whether policy significantly lowers the levels of transitory risk that households face. The results in Table 5 suggest that while policy qualitatively reduces the level of risk for the volatility measure of transitory income risk, this change is not statistically significant. In contrast, taxes and national insurance on their own, benefits on their own and benefits net of taxes and national insurance all contribute to significantly to lowering the asymmetry measure of transitory income risk.

Table 5: Policy effects on transitory income risk

	volatility	asymmetry
[1]	$[m_2^\varepsilon]^g - [m_2^\varepsilon]^{g-t-ni}$	$[m_3^\varepsilon]^g - [m_3^\varepsilon]^{g-t-ni}$
est	0.0029	-0.0105**
s.e.	(0.0107)	(0.0061)
[2]	$[m_2^\varepsilon]^g - [m_2^\varepsilon]^{g+b}$	$[m_3^\varepsilon]^g - [m_3^\varepsilon]^{g+b}$
est	0.0062	-0.0222***
s.e.	(0.0103)	(0.0053)
[3]	$[m_2^\varepsilon]^g - [m_2^\varepsilon]^{g+b-t-ni}$	$[m_3^\varepsilon]^g - [m_3^\varepsilon]^{g+b-t-ni}$
est	0.0080	-0.0274***
s.e.	(0.0099)	(0.0051)

Overall, our findings regarding the beneficial impact of social insurance policy generally are 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 US in Kniesner and Ziliak (2002), and for the US, Germany and Sweden in Busch *et al.* (2018), among others. Importantly, we find that in all cases of second and third central moments considered, for both persistent and transitory shocks, the effects of benefits in reducing risk exposure are bigger than taxes and national insurance. This is consistent with evidence from different analysis in the UK (see e.g. figure 7a Belfield *et al.* (2017) and figures 4.5 and 4.6 in Blundell and Etheridge (2010)), which suggests that benefits have stronger effects in reducing household income inequality than taxes. In contrast, Kniesner and Ziliak (2002) find that, in the US, the effects of taxes and transfers are quantitatively similar when studying the reduction in the variance of household earnings growth.

5.6 Effect of policy on the cyclicality of risk

In Table 6 we compare the cyclical behaviour of income risk (based on the third moment) pre- and post-policy to assess the effectiveness of social insurance to mitigate this risk.[16] To this end, we make row-wise comparisons of the moment estimates in the last column reported in Table 1 in columns 1 and 2 in Table 6, starting with gross income. In column [3] of this Table we formally

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

test whether the cyclical asymmetry of gross income risk is greater than the cyclical asymmetry of income risk post-policy. In other words, has counter-cyclical left-skewness been reduced by social policy? To this end, we employ a one-sided test of the null hypothesis that $H_0 : [m_3^{\eta,e} - m_3^{\eta,c}]^g \leq [m_3^{\eta,e} - m_3^{\eta,c}]^{pp}$ against the alternative hypothesis that $H_A : [m_3^{\eta,e} - m_3^{\eta,c}]^g > [m_3^{\eta,e} - m_3^{\eta,c}]^{pp}$.

Table 6: Policy effects on the cyclical asymmetry of income risk

	[1]	[2]	[3]
	$[m_3^{\eta,e} - m_3^{\eta,c}]^g$	$[m_3^{\eta,e} - m_3^{\eta,c}]^{g-t-ni}$	[1]-[2]
est	0.0324***	0.0242***	0.0082
s.e.	(0.0126)	(0.0099)	(0.0159)
	$[m_3^{\eta,e} - m_3^{\eta,c}]^g$	$[m_3^{\eta,e} - m_3^{\eta,c}]^{g+b}$	[1]-[2]
est	0.0324***	0.0072*	0.0252**
s.e.	(0.0126)	(0.0050)	(0.0132)
	$[m_3^{\eta,e} - m_3^{\eta,c}]^g$	$[m_3^{\eta,e} - m_3^{\eta,c}]^{g+b-t-ni}$	[1]-[2]
est	0.0324***	0.0051*	0.0273**
s.e.	(0.0126)	(0.0036)	(0.0132)

The results reported in column 3 in Table 6 first suggest that taxes and national insurance contributions do not significantly reduce the pro-cyclical asymmetry of gross income. In contrast, the degree of pro-cyclical asymmetry in gross income has been significantly reduced when benefits on their own are taken into account and when benefits net of taxes and national insurance are considered.[17]

These results underline the importance of benefits as a policy instrument to mitigate the increase in risk in contractions. This effect is distinct from existing results for the US, Germany and Sweden in e.g. Busch *et al.* (2018) which emphasise the importance of taxes in reducing income risk. The different results may be driven by differences in risk measures and methodological approaches employed. In particular, we study the effect of policy on the cyclicality of the skewness of the distribution of shocks to the persistent component of household income, whereas Busch *et al.* (2018) examine the effect of policy on the reducing the cyclicality of the skewness of the distribution of annual earnings growth across households. Nonetheless, as discussed

[17] To assess the robustness of our key results reported in Tables 1, 2 and 6, we use the time-series information to restrict the number of moments to estimate. In particular, we average the moments across the age groups for each period t (see Domeij and Floden (2010)). This procedure produces $2 \times T \times (T+1)/2 = 2 \times 171$ moments to match instead of $2 \times 5,187$ moments. These results are reported in Appendix D and cohere very well with our key results.

above, our results for GB cohere well with the data reported in Figure 2 and provide evidence which complements other UK findings relating to the importance of benefits in reducing income volatility and inequality in the UK (see e.g. Blundell and Etheridge (2010) and Belfield *et al.* (2017)).

6 Conclusions

Using the BHPS data from 1991-2008, this paper confirmed existing findings in the literature and established new evidence relating to the cyclical behaviour of idiosyncratic household income risk and the effect of social insurance (tax-benefits) policy in reducing this risk. State dependent persistent income risk was measured by the variance and the skewness of the probability distribution of shocks to the persistent component of idiosyncratic income in both expansions and contractions of the aggregate economy. In contrast, constant transitory income risk was measured by the variance and the skewness of the probability distribution of shocks to the transitory component of idiosyncratic income. To examine the consequences of social insurance, we estimated risk for different measures of pre- and post-policy household income and evaluated the effects of tax and benefits policies on the level and the cyclicality of risk.

Our key finding for GB is that household income risk rises in contractions implying a higher probability of receiving large negative income shocks during this state. This finding confirms, using British data, similar findings for other countries. It adds to this literature by providing evidence that in GB it is the skewness of the distribution of the shocks to the persistent component of idiosyncratic income that falls in contractions.

However, we also find that a large part of the increased risk in bad times is mitigated by social insurance policy. This effect in GB is distinct from results for the US, Germany and Sweden reported in Busch *et al.* (2018), which emphasise the importance of taxes in reducing income risk. In contrast, we find that cyclical asymmetric income risk is reduced mainly via benefits policy, confirming the importance of this instrument in mitigating income volatility and inequality previously noted by other UK studies using different methods than those employed here (see e.g. Blundell and Etheridge (2010) and Belfield *et al.* (2017)).

References

[1] Altonji, J. G., and L. M. Segal (1996). Small-sample bias in GMM estimation of covariance structures, *Journal of Business & Economic Statistics*, 14, 353-366.

[2] Bardasi, E., Jenkins, S., Sutherland, H., Levy, H., and F. Zantomio (2012). British Household Panel Survey derived current and annual net household income variables, Waves 1–18, 1991–2009. University of Essex, Institute for Social and Economic Research, Colchester, Essex: U.K. Data Archive, SN, 3909.

[3] Bayer, C. and F. Juessen (2012). The life-cycle and the business-cycle of wage risk: Cross-country comparisons, *Economic Letters*, 117, 831–833.

[4] Belfield, C., Blundell, R., Cribb, J., Hood, A., and R. Joyce (2017). Two decades of income inequality in Britain: the role of wages, household earnings and redistribution, *Economica*, 84, 157-179.

[5] Blundell, R., and B. Etheridge (2010). Consumption, income and earnings inequality in Britain, *Review of Economic Dynamics*, 13, 76–102.

[6] Brav, A., Constantinides, G., and C. Geczy (2002). Asset pricing with heterogeneous consumers and limited participation: Empirical evidence, *Journal of Political Economy*, 110, 793-824.

[7] Busch, C., Domeij, D., Guvenen, F., and R. Madera (2018). Asymmetric Business Cycle Risk and Government Insurance, *Working Paper*, University of Minnesota.

[8] Busch, C., and A. Ludwig (2016). Labor Income Risk in Germany over the Business Cycle, *Working Paper*, University of Frankfurt.

[9] Cappellari, L., and S. Jenkins (2014). Earnings and labour market volatility in Britain, with a transatlantic comparison, *Labour Economics*, 30, 201-211.

[10] Constantinides, G., and D. Duffie (1996). Asset Pricing with Heterogeneous Consumers, *The Journal of Political Economy,* 104, 219-240.

[11] Domeij, D. and M. Floden (2010). Inequality trends in Sweden 1978–2004, *Review of Economic Dynamics*, 13, 179-208.

[12] Etheridge, B. (2015). A test of the household income process using consumption and wealth data, *European Economic Review*, 78, 129-157

[13] Guvenen, F., Ozkan, S., and J. Song (2014). The Nature of Countercyclical Income Risk, *The Journal of Political Economy,* 122, 621–660.

[14] Hall, P., and J. L. Horowitz (1996). Bootstrap critical values for tests based on generalized-method-of-moments estimators, *Econometrica*, 891-916.

[15] Horowitz, J. L. (2003). Bootstrap methods for Markov processes, *Econometrica*, 71, 1049-1082.

[16] Heathcote, J., Storesletten, K., and G. L. Violante (2010). The macroeconomic implications of rising wage inequality in the United States, *Journal of Political Economy*, 118, 681-722.

[17] Krebs, T. (2007). Job Displacement Risk and the Cost of Business Cycles, *American Economic Review*, 97, 664–686.

[18] Kniesner, T. J., and J. P. Ziliak (2002). Explicit versus implicit income insurance, *Journal of Risk and Uncertainty*, 25, 5-20.

[19] Low, H., Meghir, C., and L. Pistaferri (2010). Wage risk and employment risk over the life cycle, *American Economic Review*, 100, 1432-1467.

[20] MaCurdy, T. (2007). A practitioner's approach to estimating intertemporal relationships using longitudinal data: lessons from applications in wage dynamics, *Handbook of Econometrics*, 6, 4057-4167.

[21] Mankiw, N. (1986). The equity premium and the concentration of aggregate shocks, *Journal of Financial Economics*, 17, 211-219.

[22] McKay, A. (2017). Time-varying idiosyncratic risk and aggregate consumption dynamics, *Journal of Monetary Economics*, 88, 1-14.

[23] Meghir, C., and L. Pistaferri (2004). Income variance dynamics and heterogeneity, *Econometrica*, 72, 1-32.

[24] Meghir, C., and L. Pistaferri (2011). Earnings, Consumption and life cycle choices, *Handbook of Labor Economics vol. 4*, part B, 773–854.

[25] Stewart M., and J. Swaffield (2002): Using the BHPS Wave 9 additional questions to evaluate the impact of the national minimum wage, *Oxford Bulletin of Economic and Statistics*, 64 (supplement), 633-652.

[26] Storesletten, K., Telmer, C., and A. Yaron (2004). Cyclical dynamics in idiosyncratic labor market risk, *Journal of Political Economy*, 112, 695-717.

[27] Storesletten, K., Telmer, C., and A. Yaron (2007). Asset pricing with idiosyncratic risk and overlapping generations, *Review of Economic Dynamics*, 10, 519-548.

[28] Taylor, M., Brice J., Buck N., and E. Prentice-Lane (2010). British Household Panel Survey User Manual Volume A: Introduction, Technical Report and Appendices. Colchester: University of Essex.

[29] Ziliak, J. P., Hardy, B., and C. Bollinger (2011). Earnings volatility in America: Evidence from matched CPS, *Labour Economics*, 18, 742-754.

Appendix A: Data

A.1: BHPS

The main dataset used in this paper is the British Household Panel Survey (BHPS). The BHPS is a comprehensive longitudinal study for the UK running from 1991 to 2008. As a panel data survey, the BHPS tracks individuals across households over time. In the first wave, the BHPS achieved a sample size of around 5000 households (10,000 adult interviews) or a 65% response rate. After the first wave, due to sample attrition, the sample size shrank slightly. For example, in 2000 it achieved around 4200 complete interviews or a 75% response rate (see Taylor *et al.* 2010).

Since the start of BHPS in 1991, a number of additional sub-samples have been added to the survey. For example, the European Community Household Panel Survey (ECHP) sub-sample started in 1997. It was added mainly to include respondents from Northern Ireland and a low-income sample from the UK Moreover, in 1999 two more additional boost samples, for Wales and Scotland, have been added. Since the focus is on GB, to maintain the longest possible time-series dimension in our analysis, we only use the data starting in 1991 i.e. the original panel dataset. Finally, following Blundell and Etheridge (2010), we also make use of an auxiliary dataset called "Derived Current and Annual Net Household Income Variables" compiled by Bardasi *et al.* (2012).

The BHPS contains detailed information on key magnitudes of interest for this paper. In particular, earnings, hours worked and other income. Compared to other UK panel datasets for earnings, e.g. the New Earnings Survey (NES) for the period 1975-2002 and the Annual Survey of Hours and Earnings (ASHE), for the period 1997-2015, BHPS is much smaller in the cross-sectional dimension. The obvious advantages of NES and ASHE are the accuracy and the sample size, which covers the 1% of the total working population. Additionally, the time span is large enough for time series analysis since it covers the years from 1974 to 2015. However, these datasets do not provide information relating to: (i) household physical and human capital; (ii) why individuals disappear from the survey, e.g. due to an injury, unemployment spell or move to self-employment; (iii) self-employed individuals, which are a considerable percentage of the working population (approximately 14%); and (iv) individual *annual* earnings which are only available from 1999 onwards.

In contrast to the NES and ASHE, the BHPS has information on both individual and household characteristics. Therefore, it allows the examination of compositional effects (i.e. differences between individuals and households) and thus issues relating to household insurance mechanisms. Moreover, BHPS provides important human capital variables such as educational attainment. Another, important advantage of the BHPS relates to hourly pay. As noted by Stewart and Swaffield (2002), the BHPS does not suffer from the potentially serious sample selection bias that exists in the NES. Workers earning below the pay-as-you-earn tax threshold are under-represented in the NES sample. Furthermore, BHPS also covers the self-employed, the unemployed or even those who do not participate in the labour market for any reason. Finally, it provides a consistent measure of annual earnings/incomes over the whole period at hand.

A.2: Demographic and socioeconomic variables

1. **Head and relationship to head**: For each individual in the sample, BHPS reports the relationship to the head of household in any given wave. In our analysis we focus on households whose head is married. Following Blundell and Etheridge (2010), the head of the household is defined as the oldest married (or living in partnership) male within the household.

2. **Education level**: BHPS includes information on educational attainment. For the BHPS we have used the variable wQFEDHI (where the prefix w denotes wave). To examine potential heterogeneity of earnings risk in the main text, the sample is split into degree holders and non-degree holders. The former are the individuals who hold either a Higher Degree or 1st Degree, while the latter are the individuals who hold either Higher National Certificate/Diploma or teaching qualifications or A-levels/AS level/Highers or GCSE/O level/other qualification or they have no qualifications.

A.3: Income and hours variables

1. **Labour income:** is obtained from the Derived Current and Annual Net Household Income Variables dataset (Bardasi *et al.* 2012) and is equal to total household annual labour income, wHHYRLG. Imputed values can be included in "Household total earnings" only if they do not correspond to the head of the household earnings.

2. **Gross income:** is obtained from the Derived Current and Annual Net Household Income Variables dataset (Bardasi *et al.* 2012). Gross income is equal to "Labour income", plus annual investment income, wHHYRI, plus annual private transfers income, wHHYRT, plus annual occupational pension income, wHHYRP. Private transfers income totals all receipts from other transfers (including education grants, sickness insurance, maintenance, foster allowance and payments from TU/Friendly societies, from absent family members) while occupational pension income totals all receipts from non-state pension sources. Investment income sums the estimated income from savings and investments, and all receipts from rent from property or boarders and lodgers.

3. **Gross income − taxes − NI:** is obtained from the Derived Current and Annual Net Household Income Variables dataset (Bardasi *et al.* 2012) and is equal to "Labour income" minus annual national insurance contributions, wYRNI, minus annual income tax after credits, wYRTAXNT, plus annual investment income, wHHYRI, plus annual private transfers income, wHHYRT, plus annual occupational pension income, wHHYRP.

4. **Gross income + benefits:** is obtained from the Derived Current and Annual Net Household Income Variables dataset (Bardasi *et al.* 2012) and is defined as "Gross income" plus annual social benefits income, wHHYRB. Social benefits income totals all receipts from state benefits including national insurance retirement pensions.

5. **Gross income + benefits − taxes − NI:** is obtained from the Derived Current and Annual Net Household Income Variables dataset (Bardasi *et al.* 2012) and is defined as "Gross income −taxes −NI" plus annual social benefits income, wHHYRB.

A.4: Sample selection

For all of the measures discussed below, to employ a consistent sample throughout, we use the original BHPS sample excluding the observations from the boost samples after 1997.

A.5.1: Household income

We construct households from 1991-2008 by starting with the allocation of individuals to households from BHPS and retain households with a spouse/partner relationship. The household heads must be between 23-62 years of age, report non-zero labour income and their individual earnings should be reported, not imputed and above than half of the product between the minimum legal hourly wage times 520 hours. Households comprised of a single member or those that involve cohabiting but not family-related members are discarded. Moreover, the head must not be in the military and must not have missing values for region and educational attainment. Then, we discard the observations belonging to the highest 1% of the household earnings observations in each year. For the remaining households, we only keep households who are in the sample for at least three consecutive periods.

Table A.1: Sample selection in steps

selection step	households (obs.)
1. Whole sample	130,974
2. Drop proxy & non-full interviews	128,348
3. Original sample	82,355
4. Full interview of all members in household	74,602
5. Drop if no head's educational info	73,739
6. Drop if head in military	73,662
7. Drop if head's region missing	73,638
8. Keep if more than 2 adults	48,912
9. Keep if head's earnings>threshold & living with spouse	27,304
10. Keep if heads' age $\geq$23, $\leq$62	25,794
11. Drop if top1% of household total earnings	25,545
12. Keep if present at least 3 consecutive observations	21,870
ave. N obs per wave	1,215
N of unique households	2,483
ave. obs per household	8.8

Table A.2: Summary of Selected BHPS Data (1991-2008)

Variable	mean	s.d.	min	max
Head's age	41.1	9.8	23	62
HH size	3.3	1.1	2	9
Head's earnings	31,163.5	16,590.6	1,128.9	152,725.3
labour income	46,667.1	22,140.4	1,128.9	160,989.5
gross income	48,752.5	23,480.0	1,128.9	487,313.9
gross income −taxes −NI	38,328.0	17,427.0	1128.9	463,554.8
gross income +benefits	50,553.6	22,965.3	2,246.1	488,819.7
gross income +benefits −taxes −NI	40,129.2	16,955.3	2246.1	465,060.7

Note: All monetary values are expressed in 2012 prices using the RPI deflator.
The summary statistics refer to sample selection step 12 in Table A.1.

Appendix B: Persistence and Identification

The persistence parameter ρ is identified as:

$$\rho = \left[\frac{Cov\left(\mu_{i,1,1}, \mu_{i,4,4}\right) - Cov\left(\mu_{i,1,1}, \mu_{i,3,3}\right)}{Cov\left(\mu_{i,1,1}, \mu_{i,3,3}\right) - Cov\left(\mu_{i,1,1}, \mu_{i,2,2}\right)} \right]. \tag{22}$$

This expression implies that we need *at least* 4 consecutive time periods and age groups of data to estimate ρ, otherwise the model is under-identified.

However, in order to keep the exposition tractable, for the rest of the exercise we assume that $t \in T = \{1, 2, 3\}$ and $h \in H = \{1, 2, 3\}$ and that ρ is given.[18] This additional restriction implies a total of 28 empirical moments to identify 8 parameters, m_2^{χ}, m_2^{χ}, m_2^{ε}, m_3^{ε}, $m_2^{\eta,c}$, $m_2^{\eta,e}$, $m_3^{\eta,c}$, and $m_3^{\eta,e}$. To illustrate this example, we next need to specify the function that splits the time periods into expansions and contractions, e.g.:

$$\begin{bmatrix} \text{-1:} & \text{contraction} \\ \text{0:} & \text{expansion} \\ \text{1:} & \text{contraction} \\ \text{2:} & \text{expansion} \\ \text{3:} & \text{contraction} \end{bmatrix},$$

[18]Recall in the model estimated in the main body of the paper $T = 18$ and $H = 36$.

which implies

$$\begin{bmatrix} m_2^{\eta,f(-1)} = m_2^{\eta,c} & m_3^{\eta,f(-1)} = m_3^{\eta,c} \\ m_2^{\eta,f(0)} = m_2^{\eta,e} & m_3^{\eta,f(0)} = m_3^{\eta,e} \\ m_2^{\eta,f(1)} = m_2^{\eta,c} & m_3^{\eta,f(01)} = m_3^{\eta,c} \\ m_2^{\eta,f(2)} = m_2^{\eta,e} & m_3^{\eta,f(2)} = m_3^{\eta,e} \\ m_2^{\eta,f(3)} = m_2^{\eta,c} & m_3^{\eta,f(3)} = m_3^{\eta,c} \end{bmatrix}.$$

Note that the time periods 0 and -1, i.e. past periods appear in the table above. The reason is that since an agent's income has a persistent component, then she is accumulating shocks. In turn, this means that some of the agents in the sample bring with them these past shocks, and thus, the central moments of these past shocks appear in the theoretical moments. Consequently, we have extra information which we exploit to get more accurate estimates for $m_2^{\eta,c}$, $m_2^{\eta,e}$, $m_3^{\eta,c}$, and $m_3^{\eta,e}$.

B.1: Second and third moments for transitory shocks

Using periods $t = 1, 2$ and ages $h = 1, 2$, m_2^{ε} is identified via equations (9) and (11):

$$m_2^{\varepsilon} = E\left(\mu_{i,1,1}^2\right) - \rho^{-1} Cov\left(\mu_{i,1,1}, \mu_{i,2,2}\right), \tag{23}$$

and likewise m_3^{ε} is found employing equations (10) and (12):

$$m_3^{\varepsilon} = E\left(\mu_{i,1,1}^3\right) - \rho^{-1} CoSk\left(\mu_{i,1,1}, \mu_{i,2,2}\right). \tag{24}$$

B.2: Second moments for fixed effects

Using periods $t = 1$ and ages $h = 1$, m_2^{χ} is identified via equation (9):

$$m_2^{\chi} = E(\mu_{i,1,1}^2) - m_2^{\varepsilon} - m_2^{\eta,c}, \tag{25}$$

and likewise m_3^{χ} is determined employing equation (10):

$$m_3^{\chi} = E(\mu_{i,1,1}^3) - m_3^{\varepsilon} - m_3^{\eta,c}. \tag{26}$$

Thus, we can pin down m_2^{χ} and m_3^{χ} conditional on the identification of $m_2^{\eta,c}$ and $m_3^{\eta,c}$.

B.3: Moments for innovations to the persistent component

Using equation (11) along with periods $t = 1, 2, 3$ and $h = 1, 2, 3$, $m_2^{\eta,c}$ at $t = 1$ is identified as:

$$m_2^{\eta,c} = \frac{Cov\left(\mu_{i,1,1}, \mu_{i,3,3}\right) - Cov\left(\mu_{i,1,1}, \mu_{i,2,2}\right)}{\rho\left(\rho - 1\right)}. \tag{27}$$

Likewise, using equation (12) for the same t and h, $m_3^{\eta,c}$ is given by:

$$m_3^{\eta,c} = \frac{CoSk\left(\mu_{i,1,1}, \mu_{i,3,3}\right) - CoSk\left(\mu_{i,1,1}, \mu_{i,2,2}\right)}{\rho\left(\rho-1\right)}. \tag{28}$$

Thus, having identified $m_2^{\eta,c}$ and $m_3^{\eta,c}$, we have implied m_2^{χ} and m_3^{χ} via (25) and (26) as well. Indentifying equations (27) and (28) are crucial in order to determine m_2^{χ} and m_3^{χ} which will help us pin down the rest of the parameters. Without these two conditions we cannot proceed further. That is the reason why we need *at least* 3 consecutive time periods and age groups of data to identify $m_2^{\eta,c}$ and $m_3^{\eta,c}$.

So, based on the values for m_2^{χ}, m_3^{χ}, m_3^{ε} and m_2^{ε}; $m_2^{\eta,e}$ is identified via equation (9) using $t = 2$ and $h = 1$:

$$m_2^{\eta,e} = E(\mu_{i,1,2}^2) - m_2^{\chi} - m_2^{\varepsilon}, \tag{29}$$

and likewise $m_3^{\eta,e}$ is identified employing equation (10):

$$m_3^{\eta,e} = E(\mu_{i,1,2}^3) - m_3^{\chi} - m_3^{\varepsilon}. \tag{30}$$

Finally, note that when $T = H = 3$, and conditional on assumption that ρ is given, we have 8 parameters to identify, m_2^{χ}, m_2^{χ}, m_2^{ε}, m_3^{ε}, $m_2^{\eta,c}$, $m_2^{\eta,e}$, $m_3^{\eta,c}$, and $m_3^{\eta,e}$, and a total of 28 moment conditions. However, in demonstrating identification we have used exactly 8 moments, (23)-(30), but many parameters of the statistical model are already over-identified even with the minimal requirements, i.e. $size(T) = size(H) = 3$. Clearly the parameters will be even more over-identified as T and H increase.

Appendix C: Standardised third moments

Table C.1: Cyclical household income risk

	$\widetilde{m}_3^{\eta,e}$	$\widetilde{m}_3^{\eta,c}$	$\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}$	$\widetilde{m}_3^{\varepsilon}$	$\widetilde{m}_3^{\chi}$
			labour income		
est	-1.473*	-4.381*	2.908*	-7.717*	-1.190
	[-4.762,-0.291]	[-7.997,-2.075]	[0.642,$+\infty$)	[-24.937,-3.364]	[-1.940,1.171]
			gross income		
est	-0.757	-4.620*	3.863*	-7.719*	-0.912
s.e.	[-2.255,0.888]	[-7.322,-2.834]	[1.834,$+\infty$)	[-17.830,-4.341]	[-1.532,0.188]
			gross income $-$taxes $-$NI		
est	-0.634	-4.578*	3.944*	-6.336*	-0.831
s.e.	[-2.331,0.909]	[-7.378,-2.811]	[1.878,$+\infty$)	[-12.460,-3.376]	[-1.464,0.465]
			gross income $+$benefits		
est	-0.129	-1.748*	1.619*	-3.930*	-0.552
s.e.	[-1.865,1.033]	[-3.400,-0.545]	[0.007,$+\infty$)	[-7.737,-1.700]	[-1.011,0.316]
			gross income $+$benefits $-$taxes $-$NI		
est	0.045	-1.585*	1.630	-2.497*	-0.445
s.e.	[-2.202,1.295]	[-3.196,-0.321]	[-0.230,$+\infty$)	[-5.022,-0.869]	[-0.955,0.741]

Note that the CI_{90}'s for all columns are for a two-tailed test, except the difference tests which are for a one-tailed CI_{90}. Also note that the standardised measures in this and the next two tables are defined as $\widetilde{m}_3^{\eta,e} = \frac{m_3^{\eta,e}}{\left(m_2^{\eta,e}\right)^{1.5}}$; $\widetilde{m}_3^{\eta,c} = \frac{m_3^{\eta,c}}{\left(m_2^{\eta,c}\right)^{1.5}}$; $\widetilde{m}_3^{\varepsilon} = \frac{m_3^{\varepsilon}}{\left(m_2^{\varepsilon}\right)^{1.5}}$ and $\widetilde{m}_3^{\chi} = \frac{m_3^{\chi}}{\left(m_2^{\chi}\right)^{1.5}}$.

Table C.2: Policy effects on income risk (asymmetry)

	expansions	contractions	transitory
[1]	$[\widetilde{m}_3^{\eta,e}]^g - [\widetilde{m}_3^{\eta,e}]^{g-t-ni}$	$[\widetilde{m}_3^{\eta,c}]^g - [\widetilde{m}_3^{\eta,c}]^{g-t-ni}$	$[\widetilde{m}_3^{\varepsilon}]^g - [\widetilde{m}_3^{\varepsilon}]^{g-t-ni}$
est	-0.123	-0.042	-1.383
s.e.	(-∞,1.727]	(-∞,2.484]	(-∞,3.725]
[2]	$[\widetilde{m}_3^{\eta,e}]^g - [\widetilde{m}_3^{\eta,e}]^{g+b}$	$[\widetilde{m}_3^{\eta,c}]^g - [\widetilde{m}_3^{\eta,c}]^{g+b}$	$[\widetilde{m}_3^{\varepsilon}]^g - [\widetilde{m}_3^{\varepsilon}]^{g+b}$
est	-0.628	-2.872*	-3.7891
s.e.	(-∞,1.292]	(-∞,-1.029]	(-∞,0.285]
[3]	$[\widetilde{m}_3^{\eta,e}]^g - [\widetilde{m}_3^{\eta,e}]^{g+b-t-ni}$	$[\widetilde{m}_3^{\eta,c}]^g - [\widetilde{m}_3^{\eta,c}]^{g+b-t-ni}$	$[\widetilde{m}_3^{\varepsilon}]^g - [\widetilde{m}_3^{\varepsilon}]^{g+b-t-ni}$
est	-0.802	-3.035*	-5.223*
s.e.	(-∞,1.366]	(-∞,-1.217]	(-∞,-1.864]

Table C.3: Policy effects on the cyclical asymmetry of income risk

	[1]	[2]	[3]
	$[\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}]^g$	$[\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}]^{g-t-ni}$	[1]-[2]
est	3.863*	3.944*	-0.081
s.e.	[1.834,+∞)	[1.878,+∞)	[-3.182,+∞)
	$[\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}]^g$	$[\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}]^{g+b}$	[1]-[2]
est	3.863*	1.619*	2.244
s.e.	[1.834,+∞)	[0.007,+∞)	[-0.161,+∞)
	$[\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}]^g$	$[\widetilde{m}_3^{\eta,e} - \widetilde{m}_3^{\eta,c}]^{g+b-t-ni}$	[1]-[2]
est	3.863*	1.630	2.233
s.e.	[1.834,+∞)	[-0.230,+∞)	[-0.257,+∞)

Appendix D: Robustness

The following results are based on the restricted set of 2 × 171 moments discussed in footnote 17 of the main text.

Table D.1: Cyclical household income risk

	$m_2^{\eta,e}$	$m_2^{\eta,c}$	$m_2^{\eta,e}-m_2^{\eta,c}$	$m_3^{\eta,e}$	$m_3^{\eta,c}$	$m_3^{\eta,e}-m_3^{\eta,c}$
			labour income			
est	0.0546***	0.0560***	-0.0014	-0.0155*	-0.0341***	0.0186*
s.e.	(0.0077)	(0.0093)	(0.0077)	(0.0084)	(0.0101)	(0.0133)
			gross income			
est	0.0460***	0.0498***	-0.0038	-0.0054	-0.0335***	0.0282**
s.e.	(0.0064)	(0.0087)	(0.0073)	(0.0069)	(0.0096)	(0.0123)
			gross income −taxes −NI			
est	0.0378***	0.0401***	-0.0023	-0.0030	-0.0241***	0.0210**
s.e.	(0.0047)	(0.0072)	(0.0063)	(0.0050)	(0.0075)	(0.0096)
			gross income +benefits			
est	0.0295***	0.0339***	-0.0044	0.0005	-0.0061*	0.0066*
s.e.	(0.0031)	(0.0058)	(0.0053)	(0.0031)	(0.0035)	(0.0050)
			gross income +benefits −taxes −NI			
est	0.0232***	0.0280***	-0.0048	0.0010	-0.0040*	0.0050*
s.e.	(0.0022)	(0.0044)	(0.0045)	(0.0022)	(0.0024)	(0.0034)

Table D.2: Persistence and remaining moments

	ρ	m_2^{ε}	m_3^{ε}	m_2^{χ}	m_3^{χ}
			labour income		
est	0.7846***	0.0170**	-0.0439***	0.0683***	-0.0246**
s.e.	(0.0388)	(0.0068)	(0.0058)	(0.0123)	(0.0104)
			gross income		
est	0.7907***	0.0166***	-0.0343***	0.0722***	-0.0222**
s.e.	(0.0392)	(0.0059)	(0.0050)	(0.0124)	(0.0099)
			gross income −taxes −NI		
est	0.8059***	0.0157***	-0.0242***	0.0587***	-0.0162**
s.e.	(0.0352)	(0.0045)	(0.0037)	(0.0104)	(0.0080)
			gross income + benefits		
est	0.8222***	0.0146***	-0.0129***	0.0536***	-0.0090**
s.e.	(0.0325)	(0.0037)	(0.0021)	(0.0105)	(0.0035)
			gross income +benefits −taxes −NI		
est	0.8301***	0.0131***	-0.0074***	0.0419***	-0.0057**
s.e.	(0.0275)	(0.0026)	(0.0016)	(0.0080)	(0.0024)

Table D.3: Policy effects on the cyclical asymmetry of income risk

	[1]	[2]	[3]
	$[m_3^{\eta,e} - m_3^{\eta,c}]^g$	$[m_3^{\eta,e} - m_3^{\eta,c}]^{g-t-ni}$	[1]-[2]
est	0.0282**	0.0210**	0.0071
s.e.	(0.0123)	(0.0096)	(0.0159)
	$[m_3^{\eta,e} - m_3^{\eta,c}]^g$	$[m_3^{\eta,e} - m_3^{\eta,c}]^{g+b}$	[1]-[2]
est	0.0282**	0.0066*	0.0216*
s.e.	(0.0123)	(0.0050)	(0.0133)
	$[m_3^{\eta,e} - m_3^{\eta,c}]^g$	$[m_3^{\eta,e} - m_3^{\eta,c}]^{g+b-t-ni}$	[1]-[2]
est	0.0282**	0.0050*	0.0231**
s.e.	(0.0123)	(0.0034)	(0.0128)

Figure 1: Volatility and Asymmetry of Labour Income Growth

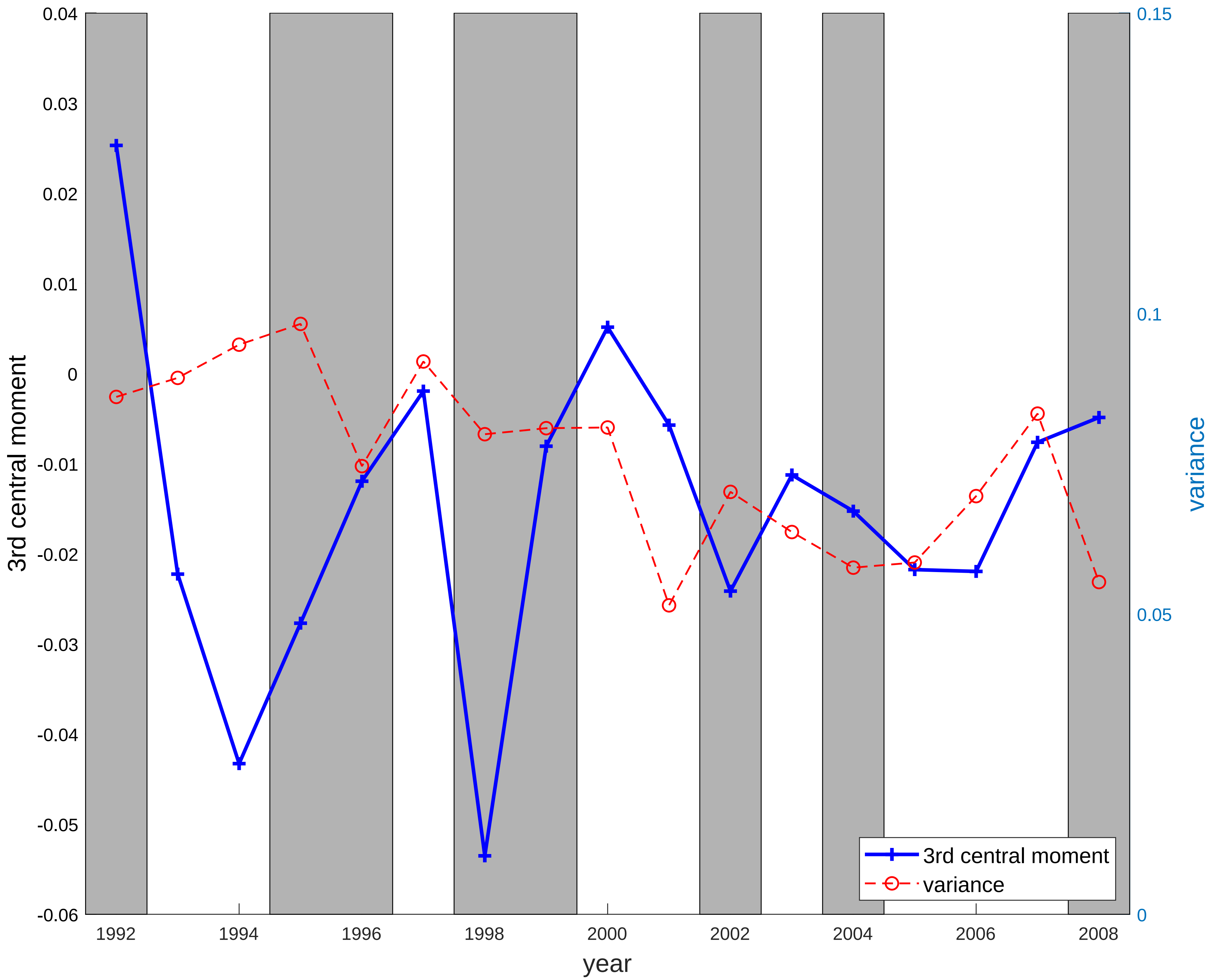

Figure 2: Tax and Benefits Effects on Asymmetry and Volatility of Income Growth