The Impact of Monetary Policy on Inequality in the UK. An Empirical Analysis*

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Abstract

The UK has experienced a dramatic increase in earnings and income inequality over the past four decades. We use detailed micro level information to construct quarterly historical measures of inequality from 1969 to 2012. We investigate whether monetary policy shocks played a role in explaining this increase in inequality. We find that contractionary monetary policy shocks lead to an increase in earnings, income and consumption inequality and contribute to their fluctuation. The response of income and consumption at different quantiles suggests that contractionary policy has a larger negative effect on low income households and those that consume the least when compared to those at the top of the distribution. Our evidence also suggests that the policy of quantitative easing may have contributed to the increase in inequality over the Great Recession.

Keywords: Inequality, Earnings, Income, SVAR, Monetary Policy Shocks
JEL No. E2, E3, E4, E5

1 Introduction

The latest financial and sovereign crises left Western economies with rising levels of inequality. A number of studies (e.g. Belfield, Cribb, Hood and Joyce, 2014; Blundell and Etheridge, 2010; Brewer and Wren-Lewis, 2012) present evidence of rising income inequality for the UK up to 2007-8. According to Belfield et al. (2014) the Gini coefficient for UK households' disposable income has increased over the last 45 years from 0.25 in 1967 to 0.36 in 2007-8. Similar trends appear for net labour earnings where the Gini increased from 0.32 in 1968 to 0.35 in 2008 (Brewer and Wren-Lewis, 2012).

A growing area of research is trying to explain the rising trend and to identify the contributing factors. Skill based education and technological advances, changes in the family structure, employment status and occupation, structural reforms in the labour market, globalisation and increased international trade have all contributed to wage and income inequality (see for example Card, 2001; Bound and Johnson, 1992; Feenstra and Hanson, 2008). However, the above factors are only part

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of the story: trying to decompose changes in income inequality, Brewer and Wren-Lewis. (2012) find that a large amount of the UK income inequality for the period 1968-2007 remains unexplained and this amount has increased to over 50% over the total variation towards the end of the period.

While fiscal policy has received substantial attention as a contributing factor to inequality, the role of monetary policy is still to be decided. Earlier studies present a contradictory view on the matter: Galbraith (1998), for example, argues that strict inflation targeting policies caused a series of recessions, higher unemployment rates and therefore a rise in inequality. Coibion et al. (2012) note that the expansionary monetary policy of Federal Reserve boosted share prices benefiting mainly shareholders, participants of financial markets and trade, who are usually the wealthier households. In addition, low income households hold most of their wealth in liquid assets which are the most vulnerable to inflation inducing monetary policies. In a recent contribution, Auclert (2016) points out that the impact of monetary policy through this channel depends on the maturity duration of assets and liabilities. In particular, expansionary monetary policy tends to provide a larger benefit to households who have negative unhedged interest rate exposure – i.e. households whose maturing liabilities exceed their maturing assets. Opposite effects have been also documented: Expansionary monetary policies and low interest rates favour borrowers who may be low income households while savers and lenders are adversely affected (Doepke and Schneider, 2006).

Hence monetary policy can have an ambiguous effect on inequality. The relationship complicates further by taking into account the sources of income of households. If policy affects wages and labour income, then households for which wage is the most important source of income will be affected by more. If monetary policy substantially alters asset prices, high income households which hold financial wealth will be highly affected.

Coibion et al. (2012) investigate whether the US monetary policy has contributed to changes in consumption and income inequality. The authors use household level data from the Consumer Expenditures Survey (CEX) since 1980 at quarterly frequency to construct their different measures of inequality and to see how these measures respond to monetary policy shocks as identified by Romer and Romer (2004). Their findings suggest that contractionary monetary policy shocks significantly increase income, consumption and wage inequality among US households.

In the present study we investigate whether monetary policy shocks have affected earnings, income and consumption inequality in the UK.[1] While our work is closely related to Coibion et al. (2012), there are a number of important distinguishing features. First, our paper uses a substantially longer quarterly time series for the inequality measures - from 1969 to 2012. This period includes a number of recessions and expansions where the Bank of England used a variety of policies, with this variation providing a stronger identification of policy shocks. Second, in addition to investigating the impact of standard monetary policy, we also examine the impact of unconventional monetary policy on inequality.[2]

[1] In a previous version of this paper we use annual data on the Gini coefficient in a mixed frequency VAR to investigate this issue. See Mumtaz and Theophilopoulou (2015).

[2] Over the past year, additional studies have applied similar methods to investigate this issue for different sets of countries. This includes Guerello (2016) for the Euro Area and Furceri et.al (2016) for developed and emerging countries who find that monetary contraction raises inequality and Inui et.al (2017) for Japan who report an unstable relationship between inequality measures and policy changes.

Using a Structural Vector Autoregression (SVAR) we find that contractionary monetary policy shocks lead to an increase in earnings, income and consumption inequality. These results remain invariant to alternative specifications of the VAR. We find that the monetary policy shock makes important contributions to historical fluctuations in the inequality measures. In order to investigate the possible factors behind the increase in inequality we estimate the SVAR using data for households at different percentiles of the distribution. Results from this exercise suggest that the contractionary monetary policy shock decreases wages and income for households at the low end of the distribution while households at the upper end are less affected. This is consistent with richer households deriving a larger proportion of their income from investments. Finally, our results also suggest that the policy of quantitative easing led to an increase in the inequality measures.

These results have important policy implications at a time when the Bank of England is contemplating the possibility of switching from unconventional to conventional monetary policy. Our results suggest that policy makers need to take redistributive effects of policy changes into account.

The rest of the paper is structured as follows: Section 2 describes the variables used, their transformations and the construction of the inequality measures. Section 3 describes the estimation of the structural VAR model and the identification scheme. Section 4 presents the main results for earnings, income and consumption, while section 5 concludes.

2 Data

We construct inequality measures for four variables: disposable income, total consumption, consumption of non durables and gross wage. The first three variables are at household level while the last one is at individual level. We draw micro data from the Family Expenditure Survey (FES) from 1969 to 2012. The FES is an annual survey which provides detailed information on demographics, income, expenditure and consumption for on average of a representative sample of 7,000 UK households per year. The households who participate on FES are asked to keep a diary with their spending of a two week period. In 2001 FES merged with the National Food Survey and became the Expenditure and Food Survey (EFS) and with the Living Costs and Food Survey (LCFS) in 2008[3]. Even though the FES has been running from 1957 there are discontinuities and small samples prior to 1968 and for this reason we start our sample from 1969. Some studies (see for example Foster, 1996, van de Ven, 2011) point out representation problems with the survey: FES tends to over represent mortgage holders, people living in the countryside, older households and under represents people living in council flats, institutions (e.g. retirement homes, military), no fixed address holders, ethnic minorities, self employed, manual workers and younger households.

The variable for disposable income is defined as weekly household income net of taxes and national insurance contributions. It is summed across all members living in the same household and it is referred throughout the text as Household's Disposable Income. After keeping only the positive values and trimming we have on average 6,900 households per year until 2006 and then the average drops to 5,600 per year. Thus, in total we have around 290,000 observations of household

[3] In 1993-94 the FES changes from a calender year to financial year (April to March) and the EFS goes back to the calender year in 2006.

income for the whole sample period. The income variable is equivalised for the family size and composition by using the modified OECD scale taking as benchmark of living standard the income of a couple without dependent children.

The variable for gross wage is Gross Personal Earnings which is the normal gross wage from any type of occupation before taxes including national insurance contributions and other deductions and bonuses. Gross wage is at the individual level, converted to weekly amounts[4]. Taking into account only positive values we have on average 7,000 observations per year or around 300,000 observations over the 43 year period. Inequality measures constructed from data on wages have the smaller measurement error as gross wage is known to households with higher precision than other forms of income, however it has the disadvantage that it is only one source of income.

We use the National Accounts definition for the total consumption variable available in FES which is the sum of housing, food, alcohol, tobacco, fuel, light and power, clothing and footwear, durable household goods, other goods, transport, vehicles and services. The non durable consumption variable excludes housing and durable household goods but includes the rest of the components of the total consumption. The two consumption variables are divided by the number of people that live in the household to construct consumption per capita.

The distributions of all four variables have been trimmed by removing the top and bottom 1%. Even though the tails of the distributions have important and complex relations with monetary policy they are likely to contain measurement errors as their inclusion causes erratic shifts in the inequality measures. Thus we follow the existing literature on this issue (see for instance Brewer and Wren-Lewis, 2012) and trim the tails by 1%.

The following macroeconomic variables are also used in the analysis below: (1) GDP per capita and in real terms (code=ABMI, ONS divided by population). (2) Inflation based on the Consumer Price Index (CPI). The CPI series is based on the seasonally adjusted harmonized index of consumer prices spliced with the retail price index excluding mortgage payments. These data are obtained from the Bank of England database. The three month treasury bill rate and the nominal effective exchange rate are obtained from Global Financial Data.

2.1 Measures of Inequality

We use two widely used measures of Inequality: the Gini coefficient of levels which is one of the most commonly used measures of inequality and takes values between 0 (perfect equality) and 1 (perfect inequality). The second measure is the cross sectional standard deviation of log levels which removes zero values thus reducing sensitivity to extreme values. Following Cloyne and Surico (2017), we assign households to different quarters within a year based on the date of survey interview. This allows us to calculate the measures of inequality at a quarterly frequency.

Figure 1 shows the evolution of the Gini coefficients for disposable income, total consumption, consumption of non durables and gross personal wage. During the 1970s and until the end of the decade a decrease in the inequality measures is observed mostly through wages as high earners experienced fall of their wages relative to low earners. Relative earnings of women increased and

[4] If the individual works full time, the weekly payment is defined as earnings, while in the case of a part time or odd job, the last payment is counted.

so did the pension benefits. The second part of 1970s is characterised by monetary easing as the nominal interest rates were kept below the actual and perspective inflation rate (Nelson, 2000).

During the period from 1979 to 1989, monetary policies aiming to control inflation were introduced. Nelson (2000) notes that nominal interest rates were persistently below the inflation rate before 1980 and persistently above after. The same period is characterised by a dramatic increase of inequality especially in disposable income. This has been attributed to higher unemployment, increase of part time work and lower working hours of the employed and higher dispersion of wages between low and high earners (Brewer and Wren-Lewis, 2012). The highest increase was that of disposable income inequality. Even though income inequality was the lowest at the beginning of the sample period, it catches up rapidly with consumption inequality in the mid 1980s. It is interesting to note that from the mid-1980s to the early 1990s inequality in non-durable consumption was lower than income inequality. This possibly reflects the impact of financial liberalization that took place in mid-1980s and enabled greater access to consumer credit. In 1992 policies of inflation targeting were introduced and in 1997 the Bank of England gained operational independence. Inequality levels for income and earnings didn't change dramatically but the Gini coefficients for consumption increased during this decade.

In the beginning of 2000s, inequality fell possibly due to a decrease of the investment income and the collapse of the dotcom bubble. In 2007 financial markets collapsed and the Great Recession followed resulting in a deep fall of inequality for all measures, especially for consumption. In 2008 the Bank of England implemented unconventional, zero bound monetary polices and Quantitative Easing. Interestingly, the Gini coefficients for consumption and earnings increase substantially after 2010 while the one for disposable income remains at low levels. Overall, during the sample period and from all the four variables, the Gini coefficient of disposable income experiences the highest rise.

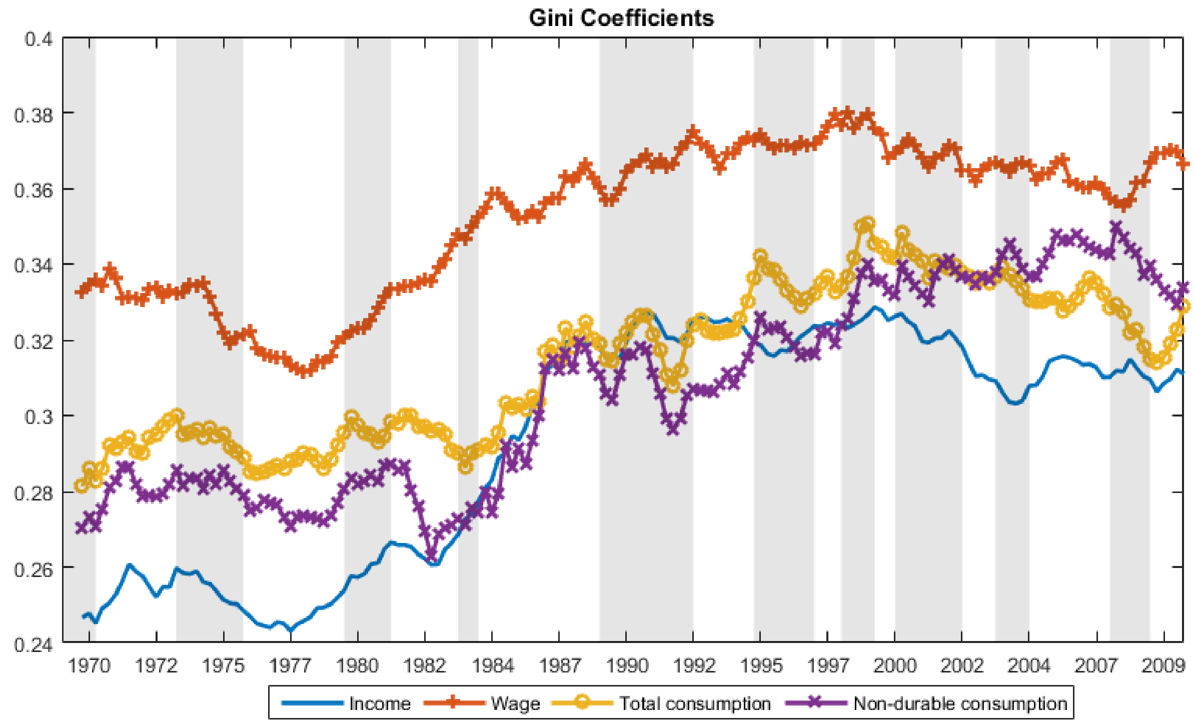

Figure 1: Gini Coefficients (4 quarter moving average). Shaded areas represent recessions as identified by the OECD.

3 Empirical Model

In order to estimate the impact of monetary policy shocks on the constructed inequality measures we use a Structural VAR model. The benchmark model is defined as

$$Z_t = c + \sum_{j=1}^{P} B_j Z_{t-j} + v_t \tag{1}$$

where $v_t \tilde{} N(0, \Omega)$. The matrix of endogenous variables includes the standard set used for small open economies: i.e. real GDP per capita, CPI inflation, the three month treasury bill rate and the growth of the nominal effective exchange rate. As discussed below, restrictions on the contemporaneous response of these variables allow us to identify a monetary policy shock. We augment this VAR model with each of the inequality measures described above, in order to estimate the impact the monetary policy shock on inequality related to earnings, income or consumption. All variables except the interest rate and the inequality measure enter in log differences. The lag length P is set to 4 in the specifications above.

We adopt a Bayesian approach to estimation and use a Gibbs sampling algorithm to approximate the posterior distribution of the model parameters. As discussed in Uhlig (2005), this approach offers a convenient method to estimate error bands for impulse responses. Note, however, that we use a flat prior and, therefore, the results reported below are data driven. The estimation algorithm is described in detail in the technical appendix to the paper.

3.1 Identification of the monetary policy shock

The covariance matrix of the residuals Ω can be decomposed as $\Omega = A_0 A_0'$ where A_0 represents the contemporaneous impact of the structural shocks ε_t:

$$v_t = A_0 \varepsilon_t \tag{2}$$

Restrictions on the elements of A_0 can be used to define the shock of interest. In the benchmark model we use sign restrictions to identify a monetary policy shock, imposing them on the appropriate elements of A_0 thus restricting the contemporaneous response of the endogenous variables to this shock. We show below that the key results are robust to using different identification schemes.

We assume that a contractionary monetary policy shock leads to a contemporaneous increase in the interest rate and the nominal exchange rate (an appreciation) and a fall in GDP growth and inflation. The response of the inequality measures is left unaffected. These restrictions on the macroeconomic variables are fairly standard and are implied by open economy models such as Lubik and Schorfheide (2006). Note that for this exercise the model is estimated using data up to 2008Q4, thus avoiding the period coinciding with unconventional monetary policy.

The identification scheme is implemented as follows: we compute the structural impact matrix, denoted A_0, via the procedure introduced by Rubio-Ramírez, Waggoner and Zha (2008). Specifically, let $\Omega = PDP'$ be the eigenvalue-eigenvector decomposition of the SVAR's covariance matrix Σ, and let $\tilde{A}_0 \equiv PD^{\frac{1}{2}}$. We draw an $N \times N$ matrix K from the $N(0, 1)$ distribution and then take the QR decomposition of K. That is, we compute Q and R such that $K = QR$. We then compute a structural impact matrix as $A_0 = \tilde{A}_0 \times Q'$ and retain it if it satisfies the sign restrictions. As pointed out in Fry and Pagan (2007), this procedure provides set identification and delivers a distribution of contemporaneous impact matrices that are admissible under the sign restrictions. In order to avoid merging the information from different admissible models, we select the A_0 matrix that is closest to the median from a given number of draws from the algorithm.

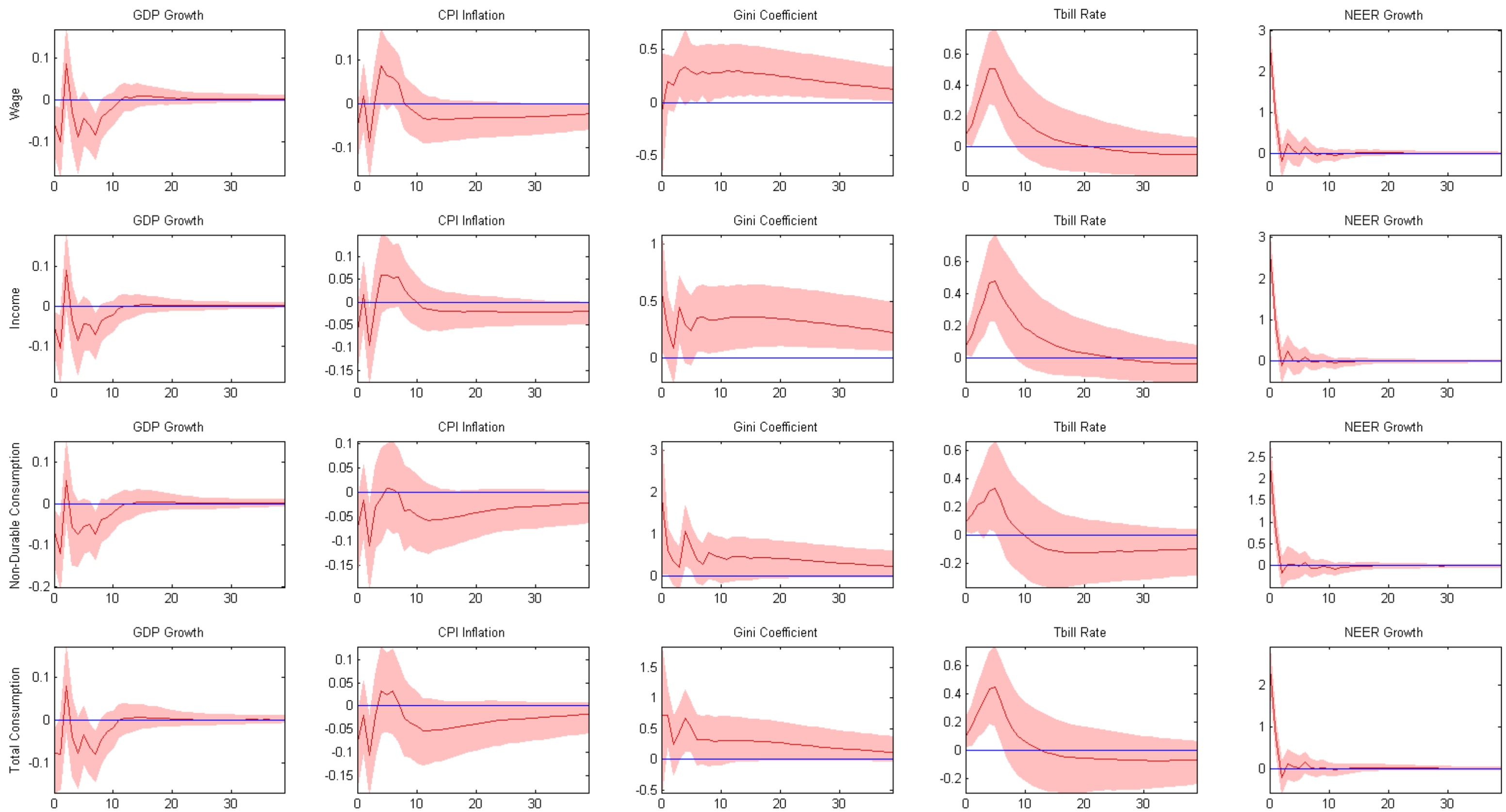

Figure 2: The impulse response of the Gini coefficient to a one standard deviation contractionary monetary policy shock. The vertical axis of each plot shows the response in percent. The red line is the median estimate and the pink shaded area is the 68% error band.

4 The response of inequality measures to monetary policy shocks

Figure 2 present the results from the benchmark VAR model. Each row shows the response to a one standard deviation contractionary monetary policy shock using the VAR model that includes the Gini coefficient derived from household data on wages, income, non-durable consumption and total consumption respectively.

The response of the macroeconomic variables to this shock is fairly standard. The 10 basis point increase in the short-term interest rate reduces GDP growth and inflation by about 0.05% to 0.1% on impact in all 4 models, while the growth of the NEER increases by about 2%-3%.

The policy shock leads to an increase in the Gini coefficient across the four models. The null hypothesis that this effect is equal to zero can be rejected in all cases. In terms of magnitude, the increase in the wage and income Gini coefficient is about 0.3% at the one year horizon. The non-durable and total consumption Gini coefficient rises by more with the response suggesting an increase of 0.7% to 1% one year after the shock. This latter result is similar to that reported by Coibion et al. (2012) for the US who find a stronger response of expenditure inequality. However, in contrast to their estimates, we find that the effect on wage and income inequality is more persistent than the impact on consumption and expenditure.[5]

The magnitude of the impulse responses suggests that the policy shock has a moderately large impact. Given the range of estimates presented above, a shock that raises the short-rate by 100 basis point is estimated to increase the Gini coefficients by 3% to 10% or by 0.009 to 0.03 in original units. To put these figures in perspective, note that the increase in the Gini coefficients observed during the 1980s amounted to about 0.06 units, i.e. a 20% change between 1980Q1 and 1990Q1.

In order to test the robustness of these results, we conduct an extensive sensitivity analysis. The detailed results from this analysis are presented in the on-line technical appendix to the paper. Here we describe the key findings.

We first consider if the results are sensitive to the assumptions regarding the identification of monetary policy shocks. The top panel of Figure A1 shows the response of the Gini coefficient from a VAR where the short-term interest rate is replaced by the Cloyne and Huertgen (2014) narrative measure of monetary policy shocks (See Appendix A). Contractionary innovations to this shock measure again lead to a persistent increase in Gini coefficients. Next, we expand the information set in the VAR. We follow the approach devised in Bernanke, Boivin and Eliasz (2005) and use a factor augmented VAR (FAVAR). The FAVAR model is estimated using a panel of 33 UK time-series spanning real activity, inflation, commodity prices and financial variables. We augment this data set with the Gini coefficients constructed from income, wages, total and non-durable consumption and identify the monetary policy shock using the benchmark sign identification scheme.[6]The bottom panels of Figure A1 shows that the key results are even stronger when incorporating a larger infor-

[5]This may reflect the fact that the persistence of wage growth was high, especially during the 1970s and the early 1980s. See for e.g. Nelson (2004).

[6]The FAVAR uses three factors. These factors explain over 70% of the variation in the dataset. The results are robust to increasing the number of factors. Details about the FAVAR model and the dataset can be found in the technical appendix.

mation set when using this alternative methodology. The four Gini coefficients show a persistent increase in the face of the contractionary policy shock with the rise in the income and wage Gini estimated to be larger than in the benchmark case. The technical appendix presents results from VARs where the monetary policy shock is identified using a Cholesky decomposition (see Figure 3 in the technical appendix). The ordering of the variables implies that monetary policy shocks have no contemporaneous impact on GDP, inflation and the Gini coefficient, but can affect the NEER immediately. The figure shows that the response of the Gini coefficient based on wage, income and total consumption expenditure is similar to the benchmark. The magnitude of the effects are marginally smaller using this identification scheme. Note, however, that the responses of inflation and the NEER are not consistent with predictions from standard open economy models–inflation rises and the exchange rate depreciates on impact in response to a monetary contraction. This potentially suggests that the recursive scheme does not identify a policy shock exclusively, leading to biased estimates of the impulse response functions. In Figure 4 in the technical appendix, we show results from VAR models where the benchmark sign restriction scheme is augmented to identify an aggregate demand and an aggregate supply shock. We use standard restrictions to identify these two shocks – a demand shock increases GDP growth, inflation and the short term rate on impact while a supply shock is restricted to move growth and inflation in opposite directions on impact. The figure shows that the main results are robust to this change in the identification scheme.

Finally, if the Gini coefficients in the VAR models are replaced with the standard deviation of logs as an alternative measure, the main results are preserved. Figure 1 in the technical appendix shows that the key results are preserved – inequality increases after a policy shock with the null hypothesis of a zero response rejected for most models.[7]

In summary, the benchmark results and the extensive sensitivity analysis provides strong evidence that contractionary monetary policy shocks are associated with an increase in inequality in the UK.

4.1 Heterogeneity of responses to policy shocks

In order to understand the possible reasons behind the response of inequality measures shown above we consider how households and individuals at different points on the distribution respond to the monetary policy shocks identified above. In particular, for each variable, we consider households and individuals that fall within the following percentiles in a given quarter: $P_1 = \left[2^{nd} : 19^{th}\right], P_2 = \left[20^{th} : 39^{th}\right], P_3 = [40^{th} : 59^{th}], P_4 = [60^{th} : 79^{th}], P_5 = [80^{th} : 98^{th}]$. We then construct measures of average real wage, real income and real per-capita consumption within these percentiles. These measures are then included in the SVAR along with four macroeconomic variables used above and their response to a monetary policy shock is examined. The shock is identified using the identification scheme discussed in section 3.1 above with no restrictions placed on the household level variables. Figure 3 plots the median response of each measure within different percentile

[7] We also consider the standard deviation of residual log income obtained after regressing household disposable income on characteristics such as age and education. The impulse response of this measure to policy shocks are very similar to the benchmark case (see technical appendix).

groups to a contractionary policy shock. The top two panels of the figure show that a monetary contraction has a large negative impact on the wages and income within the lowest percentile. Wages in P_1 decline by about 0.1% at the 2-3 year horizon, while the fall in P_2 is estimated to be about 0.05%. In contrast, the wages within the middle and higher percentiles display an initial increase before returning to base. The response of income at higher percentiles shows a similar pattern with the income in P_5 rising by 0.1%. Note, however, that in contrast, income in groups P_3, P_2 and P_1 declines. This feature matches the results for the US reported in Coibion et al. (2012) and is consistent with richer households deriving their income from financial sources. In a detailed analysis of the incomes of rich individuals, Brewer, Sibieta and Wren-Lewis (2008) show that the proportion of their income derived from investments (e.g. interest on savings) is higher than that of the average tax payer. This feature is also reflected in the FES dataset as shown in Table 1. The table shows a decomposition of gross income into shares of labour earnings and proceeds form investment and income derived from transfers. The decomposition is calculated for households in each percentile group, with P_5^* denoting the group of households within the top percentile (i.e. $P_5^* = [97^{th} : 98^{th}]$). While wages form the largest proportion of income, it is clear from the table that investment income is more important for households in the top income percentiles. It is also interesting to note that the sharp initial decline in wage and income for the group P_1 is reversed fairly quickly. For wage, this may reflect the effects of higher unemployment in the wake of the monetary contraction leading to low wage earners falling out of the distribution. As shown in Table 1, social security benefits form a larger proportion of income of this group. This may play a role in ameliorating the impact of the contractionary policy shock on income.

The bottom left panel of Figure 3 shows that the response of non-durable consumption groups is initially volatile. However, one period after the shock, expenditure rises for group P_1 while that of group P_5 declines. This pattern of responses is consistent with the possibility that households in group P_1 are credit constrained. If these households experience a persistent drop in their income as a result of contractionary monetary policy, they may substitute their expenditure from durable to non-durable goods (see Chan, Ramey and Starr (1995)). This occurs because future use of durable goods is valued less when households would like to bring the consumption of these goods to the present but are unable to do so because of credit constraints. At longer horizons, the non-durable expenditure for households in group P_1 is consistently lower than the other groups.

The bottom right panel of Figure 3 shows that inequality in total consumption expenditure is driven by the fact that all groups of households experience a persistent decline in expenditure except those in group P_5 where the fall in consumption is reversed relatively quickly and even increases for a few quarters a year after the shock. The pattern of the response for this group is similar to the response for households in top two income percentile groups. In fact, about 70% of households that fall within the top group for total expenditure have equivalized income that lies above the 60^{th} percentile. Given that investment income is more important for this group, the insights in Auclert (2016) point to the additional possibility that the households in these top percentiles have positive unhedged interest rate exposure (i.e. they might be characterised as households whose maturing assets exceed their maturing liabilities) with the interest rate increase beneficial in redistributive terms.

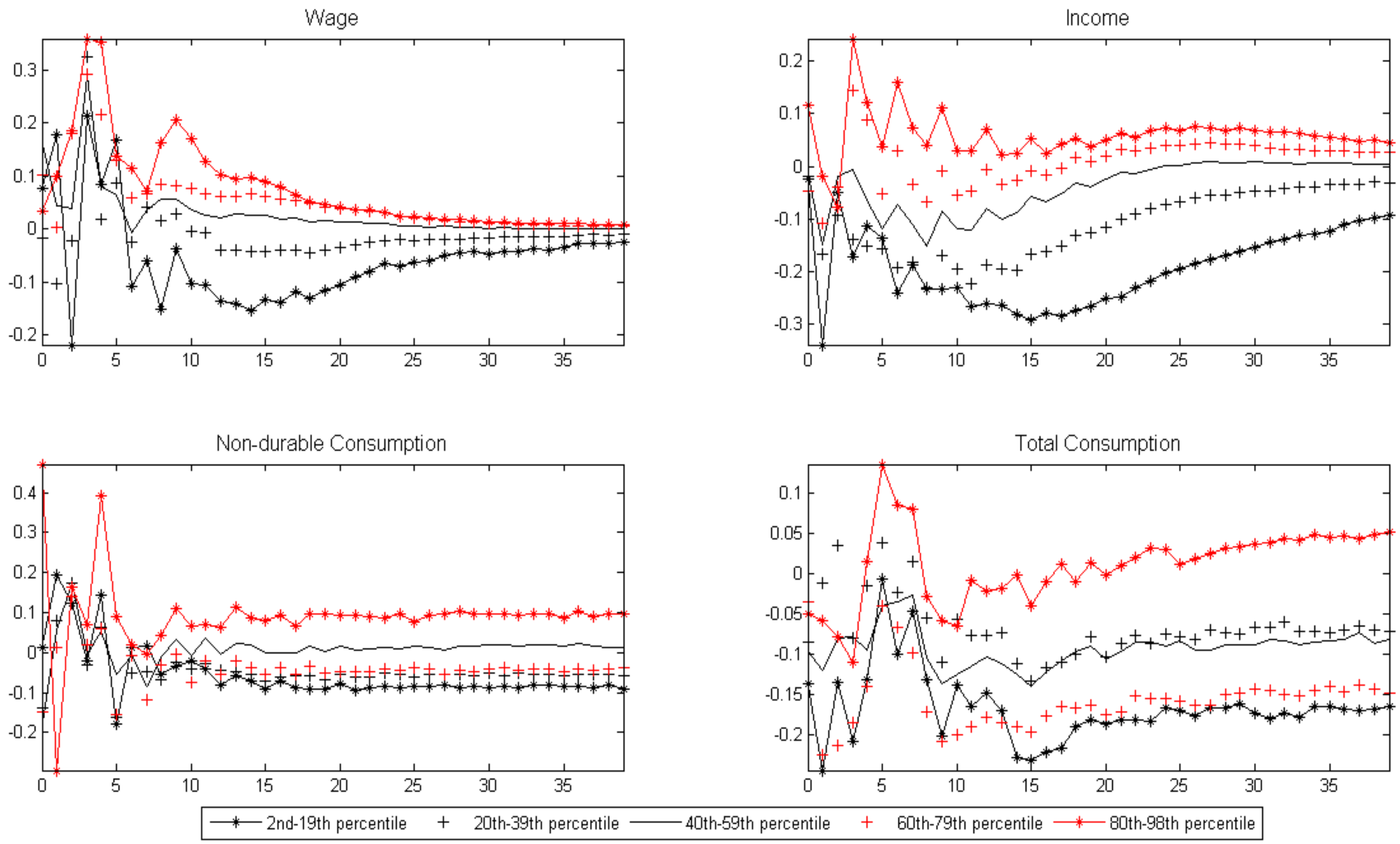

Figure 3: The response to a contractionary monetary policy shock within different percentile groups.

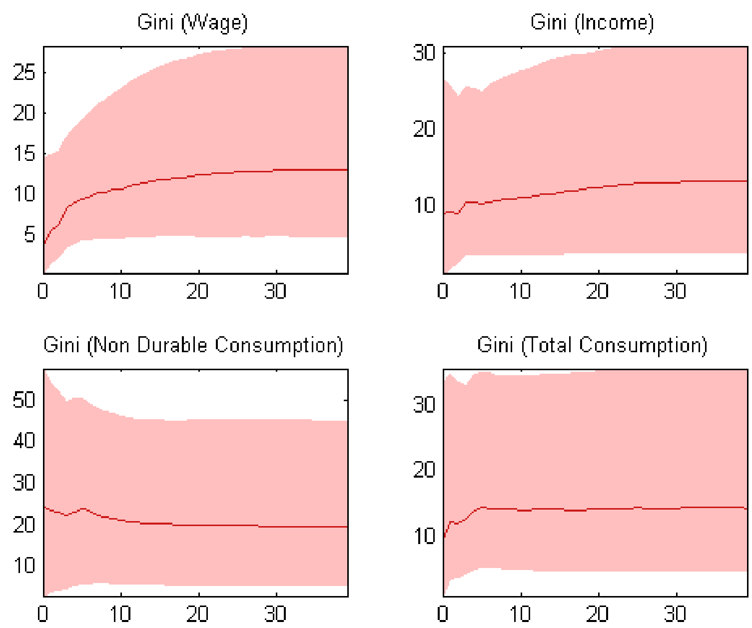

Figure 4: Percentage contribution of monetary policy shocks to the forecast error variance of the Gini coefficient. The red line is the median estimate and the pink shaded area is the 68% error band.

Quantiles	Labour Earnings	Investment Proceeds	Social Benefits
	1980s		
P_1	11.20	1.92	71.38
P_2	39.56	3.36	34.54
P_3	64.53	3.15	13.58
P_4	73.35	3.35	7.09
P_5	73.73	4.90	3.35
P_5^*	68.06	7.04	1.85
	1990s		
P_1	9.68	2.10	78.02
P_2	37.09	3.73	40.06
P_3	61.44	3.59	16.63
P_4	72.31	3.80	7.18
P_5	75.30	5.47	2.83
P_5^*	71.86	9.42	1.45
	2000s		
P_1	14.31	1.99	72.97
P_2	39.61	2.26	40.86
P_3	61.39	2.34	18.36
P_4	73.66	2.61	7.67
P_5	75.91	4.28	2.82
P_5^*	73.38	3.97	1.50
	2010-2012		
P_1	17.78	1.47	68.00
P_2	39.15	1.60	40.67
P_3	54.16	1.47	23.29
P_4	70.45	1.96	9.71
P_5	73.97	3.48	3.61
P_5^*	69.16	5.72	2.82

Table 1: Decomposition of gross income into source (percentages).

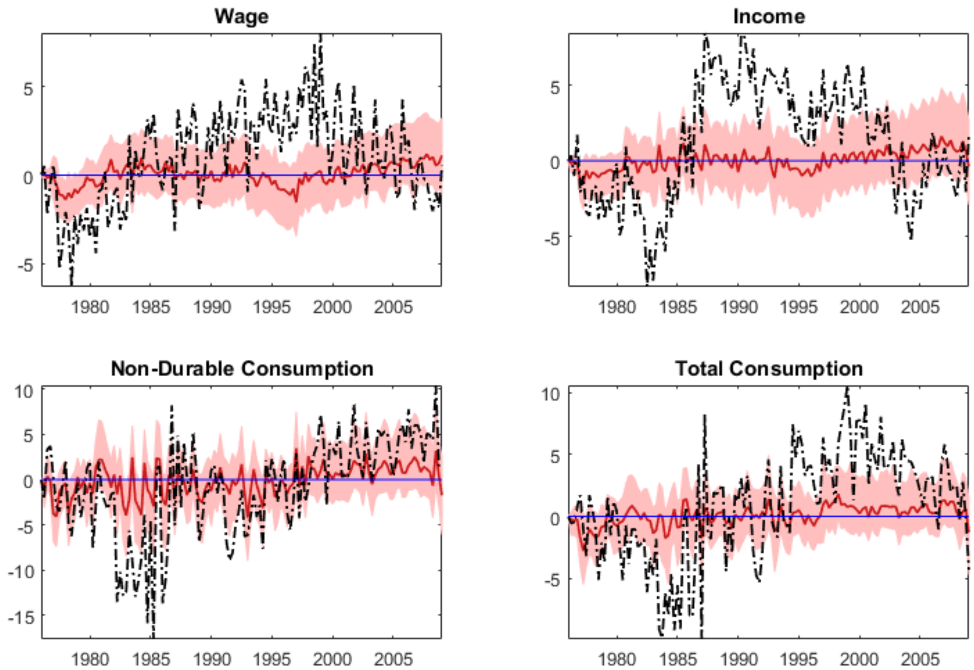

Figure 5: Historical contribution of the monetary policy shock to (de-trended) Gini Coefficient. The black dashed line shows the Gini coefficient minus its VAR implied trend. The solid red line is the median counterfactual estimate of the (de-trended) Gini coefficient assuming that only the monetary policy shock is non-zero. The pink shaded area is the 68% error band.

4.2 The contribution of monetary policy shocks to inequality

Figure 4 plots the contribution of the monetary policy shock to the forecast error variance (FEV) of the Gini coefficients. The estimated median contribution of this shock ranges from around 10% at the two year horizon for income, wage and total consumption Gini to about 20% for non-durable consumption inequality. The technical appendix shows that similar estimates are found when the standard deviation of logs is considered instead. This suggests that the policy shock made a contribution to inequality that was important both from an economic and statistical perspective. It is interesting to note that these estimates are similar to those obtained by Coibion et al. (2012) for US data.

In Figure 5, we consider if the monetary policy shock has played a role in driving cyclical fluctuations in the Gini coefficient via a historical decomposition. The contribution of the structural shocks ε_t can be calculated by noting that deviations of each variable in the VAR from a baseline trend at horizon H are given as:

$$e_t\left(H\right) = \sum_{h=0}^{H-1} \tilde{A}^h \varepsilon_{t+H-h} \tag{3}$$

where $\tilde{A}^h$ denote the impulse responses and ε_t are the shocks defined in equation 2. In Figure 5 we display this calculation for the Gini coefficient assuming that all elements of ε_t except the one corresponding to the monetary policy shock is zero. Each panel in the figure plots the four Gini coefficients in percentage deviations from the VAR implied baseline trend (black dashed line). This is compared with the counterfactual estimate of this variable under the assumption that only the monetary policy shock is operational in the VAR model (red line and shaded area). The contribution of the shock to fluctuations in the wage and income Gini are modest. It appears that policy shocks played a role during the mid and the late 1970s. During the second half of the 1970s, the UK monetary authorities cut the nominal interest rate aggressively and narrow money growth rose substantially (see Nelson, 2000). This expansionary policy appears to have made a contribution leading to a reduction in wage and income Gini over this period. A similar impact can be detected on the consumption inequality measures. It is interesting to note that after the late 1990s the median contribution of the policy shock is positive in sign for all inequality measures. One interpretation of this results is that the inflation targeting period was characterised by contractionary policy shocks as monetary authorities were more concerned with controlling inflation and this exerted upward pressure on the inequality measures.

5 The role of unconventional monetary policy

The analysis above has focussed on the role of monetary policy surprises in driving inequality. However, the post-2008 period has seen the Bank of England use Quantitative Easing (QE) or large-scale asset purchases to try and stimulate the economy. As part of this policy, the Bank purchased UK government bonds by issuing central bank money. The aim was to reduce the yield on Gilts and thus increase aggregate demand. QE can affect demand in a number of ways. However, one of the main channels of transmission is the portfolio balance channel–as the price of Gilts and related financial assets increases, demand is stimulated through wealth effects and lower borrowing costs. Note that this channel has the potential to have an asymmetric distributional impact, especially if agents at the upper end of the income or expenditure distribution have more access to the assets affected by QE.

In order to estimate the impact of QE on the measures of inequality we use two methods. First, we follow the procedure described in Kapetanios et al. (2012). This involves estimating the VAR model in equation 1 including the 10 year government bond spread as an extra variable. Thus we assume that QE affects the economy by reducing the yield on long term government bonds. Following Kapetanios et al. (2012) we assume that this reduction amounted to about 100 basis points. The VAR model is estimated up to 2008Q4. The model is then used to carry out two conditional forecasts over the next 8 quarters. The first conditional forecast assumes that the long term spread and the short-term interest rate equal their observed values over the forecast horizon. Kapetanios et al. (2012) refer to this as the 'policy scenario' as it incorporates the potential impact of QE on long term interest rates. The second conditional forecast assumes that the path of the long-term spread is higher than observed by 100 basis points over the forecast horizon while fixing the short-term interest rate at its observed value in each quarter. This simulation is assumed to represent the 'no-policy' scenario with the difference between the two conditional forecasts capturing the potential impact of QE. Of course this approach ignores any impact that QE may have through variables other than bond yields. However, given the short sample over which QE has been carried out and the subsequent difficulty of disentangling different channels, recent studies on this issue have adopted this simple approach (see Lenza, Pill and Reichlin, 2010).

As the approach of Kapetanios et al. (2012) involves out of sample forecasts from a simple VAR model, uncertainty about these projections is typically high.[8] To check the robustness of the results obtained from the Kapetanios et al. (2012) approach, we also adopt the methods used in Baumeister and Benati (2013). These authors adopt a more structural approach and use sign restrictions to identify a shock to the 10 year government bond spread in their time-varying parameter VAR model which is estimated until 2011. In particular, they assume that a contractionary spread shock leads to a contemporaneous increase in the spread, a fall in inflation and output growth but leaves the short-term interest rate unchanged. In contrast, a monetary policy shock increase the short-rate while leading to a fall in output growth, inflation and the spread. With the identified spread shock in hand, they carry out a counterfactual simulation over 2009. Under the counterfactual 'no-policy' scenario, the spread shocks are scaled up so that the simulated spread is higher than the actual values prevailing over 2009. Baumeister and Benati (2013) compare the counterfactual path of output growth with the actual path and find that the recession was estimated to be substantially worse under the counterfactual scenario. We adapt this approach for our analysis. We estimate a time-varying parameter VAR using the short-term interest rate, the 10 year government bond spread, the Gini coefficient and real GDP growth.[9] We identify a shock to the spread by assuming that it has no contemporaneous impact on the short-rate but leads to an increase the spread and a reduction in GDP growth. The monetary policy shock is assumed to increase the short-rate and reduce the spread and GDP growth contemporaneously. Following Baumeister and Benati (2013), we carry out a counterfactual experiment from 2009Q1 to 2010Q4. We scale the identified spread shock such that the counterfactual value of the spread is higher than the actual value by 100 basis points over these 8 quarters. We can then compare the implied path of the Gini coefficient with the actual path in order to investigate if the absence of QE was associated with a different outcome for the Gini coefficient.

5.1 Empirical results

In this section we present the results obtained via the counterfactual experiments based on Kapetanios et al. (2012) and Baumeister and Benati (2013). While we focus on the outcome for the Gini coefficient, the counterfactual scenarios approximating the absence of QE are associated with a peak impact of about -1% on GDP growth over the simulation period.

[8] Kapetanios et al. (2012) only present point forecasts and state that the estimates are uncertain. In our application we find that the forecast densities in both scenarios are extremely wide.

[9] Details on estimation of the model are provided in the technical appendix. We use a parsimonious 4 variable model in order to ensure that the condition of stable VAR coefficients can be satisfied at each point in time.

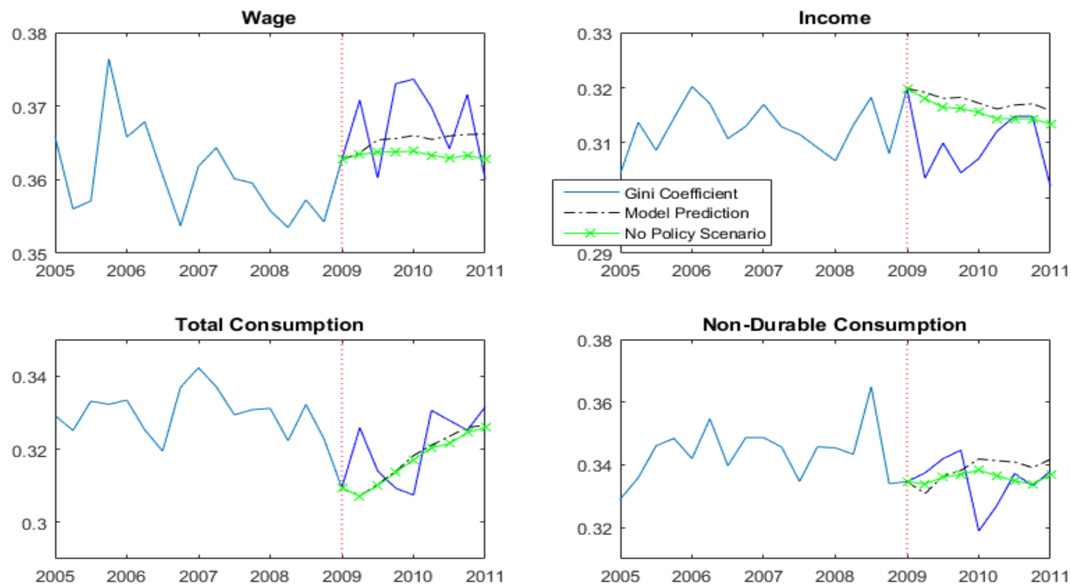

Figure 6: The impact of QE on the inequality measures

Figure 6 shows the results of the counterfactual experiment based on the Kapetanios et al. (2012) method. Each panel of the figure shows the actual data for the Gini coefficient (solid blue line), along with the median conditional forecasts under the policy scenario (black dashed line) and no-policy scenario (green starred line).

The top two panels of the figure suggest that the forecast of the wage and income Gini coefficient is higher under the policy scenario. The bottom right panel indicates that QE initially reduced inequality in consumption. However, from the second half of 2009, the non-durable consumption Gini coefficient is forecast to be lower under the no-policy scenario. It should be noted that the forecast distributions associated with these scenarios are extremely wide and indicate a high degree of uncertainty about these outcomes. We therefore consider if these results are supported by the counterfactual analysis in the spirit of Baumeister and Benati (2013).

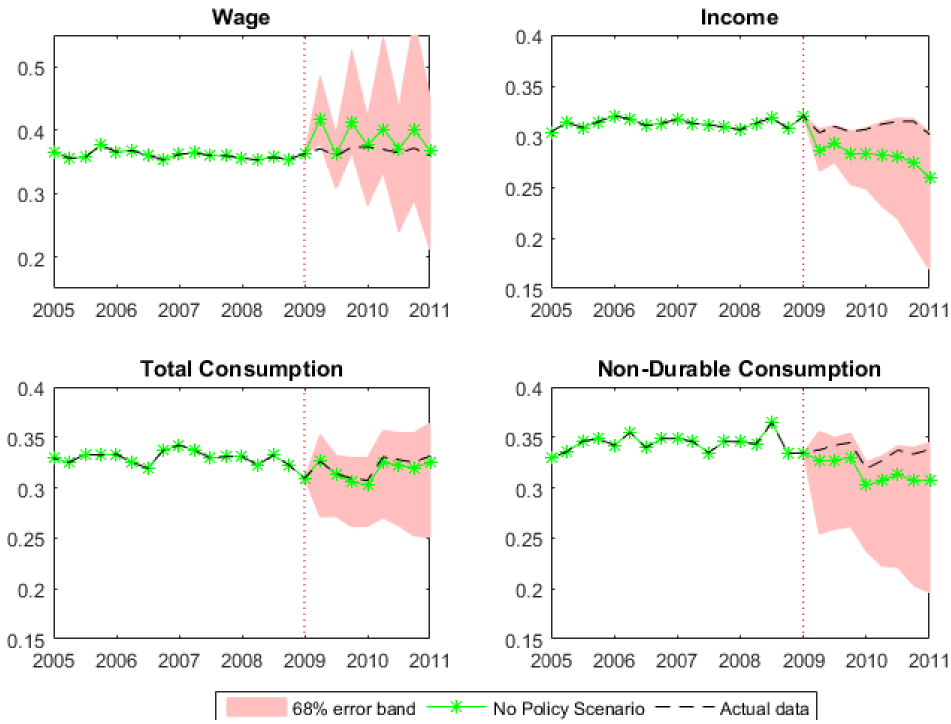

Figure 7: The impact of QE on the inequality measures using the Baumeister and Benati (2013) experiment.

The results from this counterfactual experiment are shown in Figure 7. The black dashed line represents the actual observed path of the Gini coefficient. After 2009, the green starred line is the path resulting from the counterfactual assumption that the spread was 100 basis points higher over this period. The shaded area represents the one standard deviation error band associated with the counterfactual simulation. Evidence that wage and total consumption inequality was systematically different in the absence of QE appears to be limited with the counterfactual distribution assigning a large probability to either higher or lower inequality. In contrast, the mass of the counterfactual distribution for the income and non-durable consumption Gini lies *below* the observed data suggesting some evidence for the hypothesis that QE was associated with higher income and non-durable consumption inequality.

Taking the evidence from the two counterfactual experiments together, the tentative conclusion is that QE worsened income and non-durable consumption inequality, while evidence on an impact on the remaining variables is much weaker.

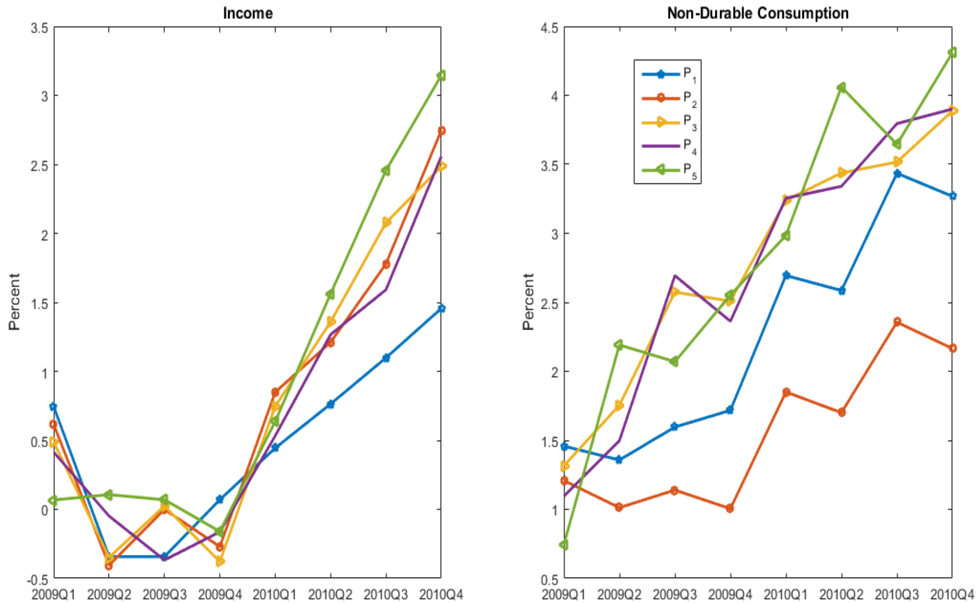

Figure 8: The impact of QE on income and consumption of percentile groups

In Figure 8, we explore this result further by re-running the first counterfactual (see Figure 6) replacing the Gini coefficients with percentile groups for income and non-durable consumption defined above.[10] The figure presents the difference in the forecasts obtained under the policy and non-policy scenario. Positive values for this quantity thus indicate that the forecast of income and non-durable consumption is higher assuming that QE depressed long-term yields by 100 basis points. The left panel of the figure shows that QE had a positive impact on the income of all groups after 2009Q3. However, it is clear from the figure that group P_5 derives the largest benefit, while the smallest increase in income occurs for group P_1. Similarly, the right panel of figure 8 shows that from 2009Q2 onwards, non-durable consumption for groups P_3, P_4 and P_5 appears have been consistently higher. These redistributive effects of QE may be driven by its impact on asset prices. QE provides a signal to the market that short-term interest rates are likely to stay low in the future. Moreover, Joyce et.al (2010) show that the transmission of this policy into long term interest rates was driven through reductions in term premia with investors more willing to hold longer term bonds thus reducing their unhedged interest rate exposure (see Auclert (2016)). If this is the case for the high income and consumption groups in our data set, QE may benefit them in a disproportionate manner.

[10] Including a large number of endogenous variables in a time-varying VAR makes estimation computationally infeasible as it is difficult to satisfy the VAR stability condition. We therefore focus on the first counterfactual which involves a fixed coefficient VAR.

6 Conclusions

This paper examines the impact of monetary policy shocks on earnings, income and consumption inequality in the UK. We build quarterly historical time series for measures of labour earnings, income and consumption inequality from the FES database. We then include these measures in a structural VAR model and estimate that contractionary monetary policy shocks lead to an increase in inequality–a one standard deviation policy contraction raises the Gini coefficient by about 1%. Impulse responses of earnings, income and consumption at the lower and upper tails of the distribution suggest that contractionary monetary policy has a larger adverse effect on these variables for the former group.

Monetary policy shocks explain a significant proportion of the fluctuations in the inequality measures with the contribution to the variance estimated to be about 20%. Historical decomposition suggests that policy shocks contributed to a decrease in inequality in mid and the late 1970s but had a positive impact over the post-inflation targeting period. Finally, we estimate that QE may have led to an increase in income inequality. Households who hold financial assets which experienced price appreciation may have benefited more than poorer households who don't have access to financial markets. These results suggest that the beneficial macroeconomic impact of QE documented in recent papers (Kapetanios et al. (2012) and Baumeister and Benati (2013)) may have to be qualified by acknowledging the possibility of adverse distributional effects.

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A Robustness of VAR identification

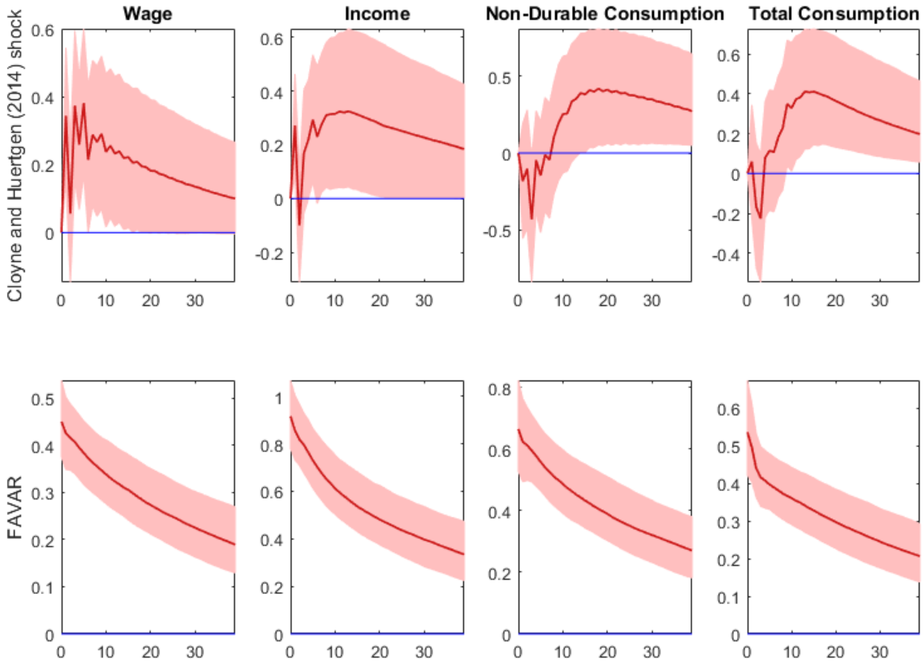

Figure A1: Response of Gini coefficients from VARs using narrative measure of monetary policy shock (top rows) and a FAVAR (bottom rows)