Wage Led Aggregate Demand in the United Kingdom*

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Abstract

The wage led aggregate demand hypothesis is examined for the United Kingdom over the period 1971 - 2007. Existing studies disagree on the aggregate demand regime for the UK, and this appears to be due to differing empirical approaches. Studies relying on equation-by-equation estimation procedures with annual data tend to find support for wage led aggregate demand in the UK, whilst studies using systems estimation procedures or quarterly data tend to find less or no support for the hypothesis. The present paper attempts to resolve this incongruity by testing the wage led aggregate demand hypothesis for the UK with structural VAR models estimated on quarterly data. In addition, the paper focuses on movements in the real wage, rather than the labour share, as this allows us to take a relatively ambivalent approach to identification. Given this, we estimate models in GDP, total employment, and the real wage, which provide robust evidence in favour of the wage led aggregate demand hypothesis.

Keywords: Real Wages, Income Distribution, Business Cycles.

JEL Codes: E32, E25, B50, E12.

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*We would like to thank . . .

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1 Introduction

The wage led aggregate demand hypothesis can be traced back to the work of Michał Kalecki and Josef Steindl via Rowthorn (1981), Dutt (1984), and Bhaduri and Marglin (1990). The theoretical framework supposes that total consumption is an increasing function of the labour share, and total investment and net exports are decreasing functions of the labour share. Depending on the relative strength of these distributive effects, an increase in the labour share will either increase or decrease aggregate demand, and thereby gross domestic product (GDP). The wage led aggregate demand hypothesis states that the distributive effect on demand works in favour of wages: an increase in the labour share will lead to an increase in GDP.

The interplay between the distribution of income and the level of aggregate demand has been argued, by a number of prominent economists, to be of central importance in rethinking macroeconomic theory in light of the 2008 crisis (Stiglitz 2011). The wage led aggregate demand hypothesis itself has been influential in those policy-focused international institutions broadly aligned with developing countries and the labour movement (UNCTAD 2010, Lavoie and Stockhammer 2012). However, a variety of estimation procedures and data types have been used to test the hypothesis, and different studies often come to different conclusions. This is particularly the case in the UK. Studies relying on equation-by-equation estimation procedures with annual data, for example, tend to find support for wage led aggregate demand in the UK. This is the case in Bowles and Boyer (1995), Naastepad and Storm (2007), and Hein and Vogel (2008). Studies using systems estimation procedures or quarterly data tend to find less or no support for the hypothesis. This is the case in Stockhammer and Onaran (2004), which finds no support for the hypothesis, and Stockhammer and Stehrer (2011), which finds that aggregate demand in the UK is profit led.

The major purpose of the present paper is to attempt to resolve (or at least start to resolve) this incongruity, by testing the wage led aggregate demand hypothesis for the UK with structural VAR models estimated on quarterly data. Given this, our methodological approach is to test the hypothesis from a variety of angles. The first aspect of this is to focus on identified shocks to the real wage, rather than the labour share or total labour income. Modelling the labour share directly almost forces the researcher to use identification restrictions based on the Bhaduri-Marglin model, or some adaptation of it, which is the approach taken in the bulk of the existing literature. Focusing on the real wage allows us to take a fairly ambivalent approach to identification, so we do not have to specify a particular structural model with a large number of degrees of freedom. As we also use a number of data definitions and filtering techniques, this should lend our results a degree of reliability. Essentially, instead of identifying shocks to the labour share directly, we estimate shocks to the labour share indirectly, by estimating the responses of GDP and total employment to identified shocks to the real wage.

Given our use of a number of filtering techniques, and a desire to use as liberal an approach to identification as possible, we focus solely on the short run effects of changes in the functional income distribution in this paper. We do not, therefore, look for long run relationships in our data, although given the secular decrease in the UK labour share since the 1970s it is not obvious *a priori* that GDP, total employment, and the real wage should be cointegrated. This is, perhaps, the simplest interpretation of the wage led aggregate demand hypothesis, although it runs counter to arguments made in Blecker (2014). Whilst we are not averse to the proposition that income distribution effects operate at low frequencies, it seems likely that the best way to approach this problem empirically would be via cross-country growth regressions, rather than single country time series models. The foregoing, in addition, explains why we refer to the wage led aggregate demand hypothesis throughout the paper, rather than the wage led growth hypothesis as in Bhaduri (2008).

The motivation for this paper is thus to strengthen the existing literature on wage led aggregate demand in the United Kingdom. The literature disagrees on the aggregate demand regime depending on the empirical method chosen, and the identification methods used are relatively restrictive. We estimate SVARs on quarterly data, and take a relatively ambivalent approach to identification, data type, and filtering method. This should, it is hoped, increase the reliability of our estimates. Our main contribution is then to strengthen the conclusion of the bulk of the literature: aggregate demand in the UK appears to be wage led, and an unexpected increase in the labour share should be expansionary in the short run. The rest of the paper is organised as follows. Section 2 discusses the wage led aggregate demand hypothesis in the context of movements in the real wage, rather than the labour share, and compares this to the New Keynesian macro-labour approach. With this theoretical background, section 3 discusses our empirical approach and identification strategy. Section 4 discusses our data sources and variable definitions, and section 5 presents the estimation results. Section 6 concludes, and suggests avenues for future research.

2 The Wage Led Aggregate Demand Hypothesis

Consider a general model of the business cycle, where the endogenous variables are consumption (C), investment (I), government expenditure (G), net exports (X), GDP (Y), total employment (L), the labour force (N), the capital stock (K), the real wage (w), and the rate of profit (r). In order to simplify the analysis, we will suppress the capital stock and labour force under the assumption that neither variable fluctuates significantly at business cycle frequencies. In addition, we will suppress the rate of profit by assuming that it is not a behavioural variable, and is instead given by the identity,

$$r_t = \frac{(Y_t - w_t L_t)}{K}. \tag{1}$$

In this case, we can write the behavioural equations in the general form,

$$C_t = c(w_t, Y_t, L_t, Z_t), \tag{2}$$

$$I_t = i(w_t, Y_t, L_t, Z_t), \tag{3}$$

$$G_t = g(w_t, Y_t, L_t, Z_t), \tag{4}$$

$$X_t = x(w_t, Y_t, L_t, Z_t), \tag{5}$$

$$L_t = l(w_t, Y_t, Z_t), \tag{6}$$

$$w_t = h(Y_t, L_t, Z_t), \tag{7}$$

where Z_t denotes a vector of predetermined variables. Clearly, although we do not wish to specify the behavioural equations any further than is necessary, the theoretical framework underpinning (2) - (7) is largely Keynesian, with Post Keynesian and/or Marxian priors concerning the importance of the functional distribution of income.

A popular approach in the wage led aggregate demand literature, which goes back at least to Bhaduri and Marglin (1990), is to recast (2) - (7) in terms of the labour share rather than the real wage. This puts emphasis on the idea that it is the distributive mechanism that the wage led aggregate demand hypothesis is concerned with. Precisely, whilst the response of consumption to increases in the labour share can be expected to be positive, and the response of investment and net exports to increases in the labour share can be expected to be negative, the wage led aggregate demand hypothesis states that the positive effect on consumption will outweigh the negative effects on investment and net exports such that the total effect on GDP of an increase in the labour share will be positive.

This is a compelling hypothesis, as it implies that there is no trade-off between equity and growth, at least in the short run. To test the hypothesis, however, we will retain (2) - (7) as above, and focus on movements in the real wage rather than the labour share. Our rationale is that a focus on the labour share itself almost forces the researcher to use identification restrictions based on the Bhaduri-Marglin model, or some adaptation of it. This is because models in the tradition of Bhaduri and Marglin (1990) are, to the best of our knowledge, the only class of model specified in terms of the labour share. As we want to take as liberal an approach to identification as possible, we therefore specify the model in terms of the real wage. In addition, it is worth noting that the government has no direct control over the labour share, which is of direct interest in the construction of wage led aggregate demand management policy. This is, in fact, reflected in the policies proposed in the wage led growth literature, for example:

> "Pro-labour policies . . . are often referred to as policies that strengthen the welfare state, labour market institutions, labour unions, and the ability to engage in collective bargaining (e.g., by extending the reach of bargaining agreements to non-unionised firms). Pro-labour policies are also associated with increased unemployment benefits, higher minimum wages and a higher minimum wage relative to the median wage, as well as reductions in wage and salary dispersion" (Lavoie and Stockhammer 2012).

The authors go on to state that, "all else equal, with a pro-labour distributional policy, the wage share will remain constant or will increase over the long run" (*ibid.*). Thus wage led aggregate demand policies are concerned with targeting the real wage, in order to increase the labour share, using instruments such as minimum wage increases.

Given the above, in order to simplify our analysis we will not model the components of aggregate demand separately. This is because a 6 variable VAR limits the amount of lags that can be included, and makes identification rather difficult (or at least contentious, given the large number of possible identification strategies). Thus we take a linear approximation to the aggregate demand equation that solves,

$$Y_t = c(\bullet) + i(\bullet) + g(\bullet) + x(\bullet), \tag{8}$$

with $c(\bullet)$, $i(\bullet)$, $g(\bullet)$, and $x(\bullet)$ given by (2) - (5). We are then left with the following model in GDP, total employment, and the real wage:

$$Y_t = y(w_t, L_t, Z_t), \tag{9}$$

$$L_t = l(w_t, Y_t, Z_t), \tag{10}$$

$$w_t = h(Y_t, L_t, Z_t). \tag{11}$$

Here, (9) is the aggregate demand curve, (10) is a labour demand curve, and (11) is a wage curve, based on the imperfect competition approach to the labour market and/or efficiency wage theory[1]. Whilst we expect the wage curve to be increasing in both GDP and employment, both $\partial y/\partial w_t$ and $\partial l/\partial w_t$ are ambiguous. If the former is positive, then a positive shock to the wage curve increases aggregate demand. The general case is illustrated in the left panel of figure 1, which plots an upward sloping wage curve and upward sloping aggregate demand curve in (w_t, Y_t) space.

This approach, in which SVARs are estimated in GDP, total employment, and the real wage, is similar to certain macro-labour studies in the New Keynesian tradition, particularly those empirical studies in the tradition of Layard et al (1991). Balmaseda et al (2000), for example, estimates SVARs in GDP, the real wage, and the unemployment rate, where the underlying structural model includes the following equations in the real wage and GDP:

$$\omega - p = \theta, \tag{B1}$$

$$q = \varnothing(d - p) + a\theta. \tag{B2}$$

Here, ω and p denote the log nominal wage rate and output price, hence the real wage is given by the stochastic labour productivity process θ. Log output q is a function of a stochastic nominal demand process d, the price level, and labour productivity, where the latter is assumed to enter through permanent income effects on consumption (Balmaseda et al 2000: 5). Thus $a > 0$, and a positive shock to the real wage curve is expected to increase aggregate demand, but *not* through income distribution effects. In fact, the authors find that positive shocks to the wage curve do indeed increase aggregate demand, using data for a number of countries including the UK. A more recent paper with similar results is Duarte and Robalo Marques (2013).

Thus Balmaseda et al (2000) find that positive shocks to the real wage curve increase aggregate demand. However, they interpret these shocks as productivity shocks, and assume

[1]See Layard et al (2005) or Layard et al (1991) for an exposition of the imperfect competition approach to the labour market, and Akerlof and Yellen (1986) for the key efficiency wage papers.

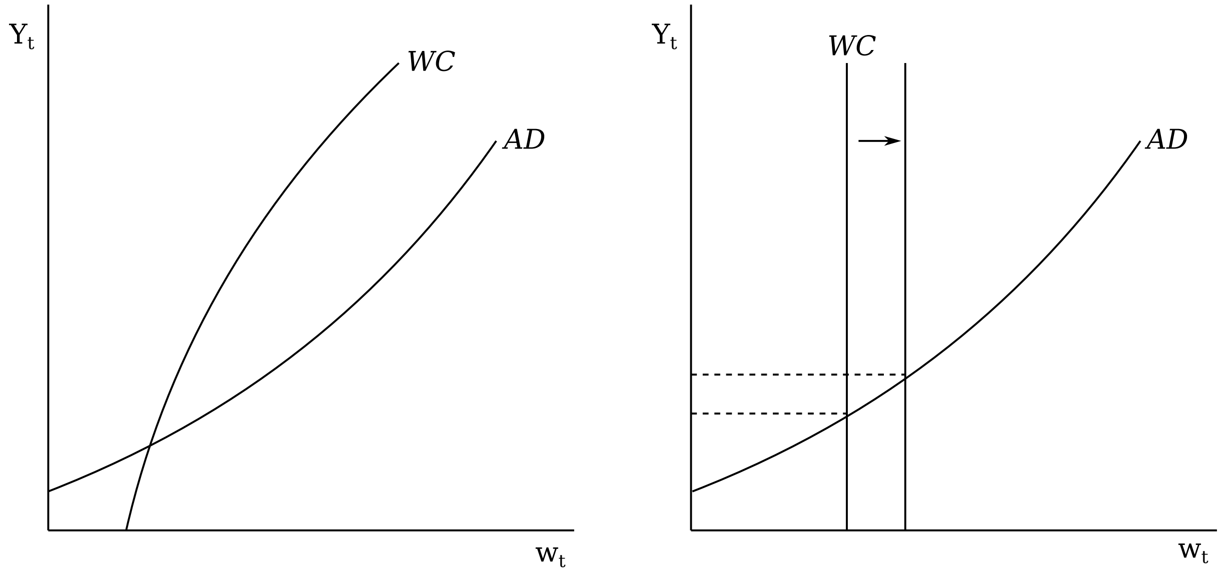

Figure 1: Illustrating the Wage Led Aggregate Demand Hypothesis.

that these shocks increase consumption directly via permanent income effects, rather than indirectly via income distribution effects. There is, therefore, a degree of observational equivalence between the wage led aggregate demand hypothesis and the permanent income interpretation of wage curve shocks in the New Keynesian macro-labour literature.

Whilst issues of observational equivalence are hard to overcome, we will attempt to alleviate the problem in this particular instance by using our identified real wage shocks to indirectly identify shocks to the labour share. Precisely, we will consider the wage led aggregate demand hypothesis to be supported if positive shocks to the wage curve are associated with *both* increases in aggregate demand *and* increases in the labour share. This is the case if the percentage increase in the real wage plus the percentage change in total employment is greater than the percentage increase in GDP. This case is illustrated in the right panel of figure 1, under the assumption that the change in total employment is non-negative. To clarify the mechanism, $\partial h / \partial Y_t = 0$ in the figure, although this will also be part of our identification strategy described in section 3.

To summarise, we approach the wage led aggregate demand hypothesis by means of a simple model in GDP, total employment, and the real wage. To test the wage led aggregate demand hypothesis, we require identified shocks to the wage curve. If positive shocks lead to increases in GDP, *and* the percentage increase in the real wage plus the percentage change in total employment is greater than the percentage increase in GDP, then the wage led aggregate demand hypothesis is supported by our analysis. Section 3 describes our empirical approach, and in particular our identification strategy.

3 Empirical Approach

The first empirical study of the wage led aggregate demand hypothesis is, to the best of our knowledge, Bowles and Boyer (1995), with subsequent studies including Stockhammer and Onaran (2004), Ederer and Stockhammer (2007), Naastepad and Storm (2007), Hein and Vogel (2008), Stockhammer et al (2009), Stockhammer and Stehrer (2011), Stockhammer et al (2011), Onaran et al (2011), Onaran and Galanis (2013), and Hartwig (2014). Studies specifically concerned with developing countries include Onaran and Stockhammer (2005), Wang (2009), Jetin and Kurt (2011), and Molero-Simarro (2015). The majority of these studies estimate linear models of the components of aggregate demand and the labour share (or total labour income and total profit income). Whilst Stockhammer and Onaran (2004) estimate a structural VAR model, the vast majority of the literature uses an equation-by-equation estimation procedure. In principle this results in estimation bias, which is usually explicitly considered and often avoided by means of lag exclusions (e.g. Stockhammer and Stehrer 2011).

Allowing for the issue of bias, the equation-by-equation procedure appears to be favoured when modelling the components of aggregate demand separately. As noted above, such a model is of particularly high dimension and would be difficult to identify in a VAR, and this problem is compounded as most of the literature uses annual or semi-annual data. As a result, the method pursued by the majority of the existing literature is closer to that of the simultaneous equation model (SEM) approach, where contemporaneous multipliers are calculated directly from coefficient estimates[2]. The main problem here, noted in Stockhammer et al (2009), Stockhammer and Stehrer (2011), and Onaran and Galanis (2013), is that the effects of interaction at various lag lengths have to be ignored.

Given this, different studies disagree on the aggregate demand regime for different countries, but most papers using the equation-by-equation method agree that aggregate demand in the UK is wage led. This is the case in Bowles and Boyer (1995), Naastepad and Storm (2007), and Hein and Vogel (2008). However, the one paper that uses the SVAR methodology (Stockhammer and Onaran 2004) finds little support for the wage led aggregate demand hypothesis, and the one paper that uses quarterly data (Stockhammer and Stehrer 2011) finds that aggregate demand in the UK is profit led. Thus there is disagreement over whether aggregate demand in the UK is wage or profit led, and this disagreement appears to stem from the differing data sources and empirical methods used.

The present paper estimates structural VAR models on quarterly UK data. We aim, therefore, to contribute to the empirical evidence on distribution effects on aggregate demand in the UK, which differ depending on the data type and empirical method chosen. As discussed in section 2, we model movements in the real wage rather than the labour share, and follow Stockhammer and Onaran (2004) by estimating the following,

$$Az_t = \alpha + \sum_{i=1}^{p} A_i z_{t-i} + Bu_t, \tag{12}$$

$$z_t = \mu + \sum_{i=1}^{p} C_i z_{t-i} + \epsilon_t, \tag{13}$$

where (12) is the structural model, and (13) is the reduced form. The vector z_t contains GDP, total employment, and the real wage, u_t is a white noise vector process with $u_t \sim N(0, I)$, and ϵ_t is a white noise vector process with $\epsilon_t \sim N(0, \Sigma)$. This leads to an identification problem: recovering A from the estimated reduced form (13). There are a number of identification strategies in the SVAR literature, with the three major approaches relying on short run restrictions, long run restrictions, and sign restrictions. The short run approach restricts

[2]See Bårdsen et al (2005) for a rigorous exposition and defence of the SEM approach to applied macroeconomics.

elements of A to equal zero, which effectively specifies zero elasticities within the period. The long run approach is more involved, but essentially involves limiting the cumulated effects of certain shocks to equal zero. If the elements of z are in differences, this is the same as imposing a zero effect on the level of the variable in question. Finally, imposing sign restrictions is more involved still, and relies on inequality constraints on impulse response functions.

Our identification strategy follows the short run restriction approach described in Sims (1986) by specifying zero elasticities within the period. The choice is simple, as we are not dealing with a simple AD-AS system in which long run restrictions make a degree of sense, and the point of the exercise is to uncover the signs of impulse response functions, so sign restrictions are of limited use. Given this, from (12) and (13) we have $\epsilon_t = A^{-1}Bu_t$, or $A\epsilon_t = Bu_t$. Expanding, we have:

$$\begin{pmatrix} a_{11} & a_{12} & a_{13} \\ a_{21} & a_{22} & a_{23} \\ a_{31} & a_{32} & a_{33} \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} b_{11} & b_{12} & b_{13} \\ b_{21} & b_{22} & b_{23} \\ b_{31} & b_{32} & b_{33} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}. \tag{14}$$

In order to normalise the system, we restrict the diagonal elements of A to equal 1, and we assume B is a diagonal matrix (i.e. each structural shock only enters a single equation directly). The shock processes are therefore interpreted as composite shocks to the variables in question, rather than underlying shock processes that can affect more than one variable directly. Our interpretation is as follows: u_Y subsumes the major sources of aggregate demand disturbance, including fiscal and monetary policy shocks; u_L subsumes the major sources of labour demand disturbance, mainly labour productivity shocks; u_w subsumes the major non-productivity sources of wage curve disturbance, mainly bargaining power shocks. For this interpretation to hold, the various parties to wage bargaining must not be able to observe labour productivity shocks within the period, as the assumption that $b_{32} = 0$ is difficult to justify otherwise.

Given the above, our identification strategy relies on two restrictions to A. Specifically, we assume $a_{31} = a_{32} = 0$. It is well established that nominal wage and price contracts are updated infrequently and set in advance of production. In addition, the relevant output and employment data are released with a lag, and are subject to substantial uncertainty concerning future revision. Therefore it seems reasonable to assume that the real wage does not react to aggregate demand or labour demand movements within the period, which supplies us with two identification restrictions, $\partial h/\partial Y_t = \partial h/\partial L_t = 0$. Thus (14) reduces to,

$$\begin{pmatrix} 1 & a_{12} & a_{13} \\ a_{21} & 1 & a_{23} \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} b_{11} & 0 & 0 \\ 0 & b_{22} & 0 \\ 0 & 0 & b_{33} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}. \tag{15}$$

Given this, we still require one additional restriction to fully identify the model. In principle we have four possible identification strategies: specifying $a_{12} = 0$, $a_{13} = 0$, $a_{21} = 0$, or $a_{23} = 0$, alongside the restrictions already made in (15). However, after experimentation, we have concluded that specifying $a_{13} = 0$ in (15) results in a very badly identified system,

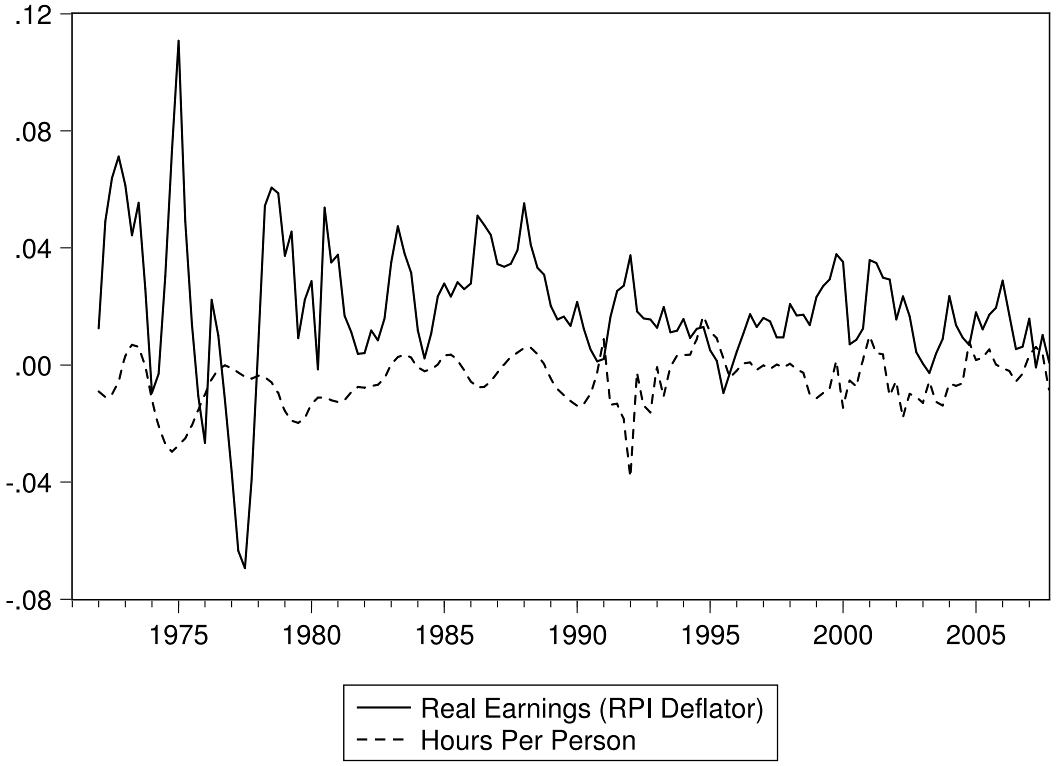

Figure 2: Real Earnings and Hours, 1972 - 2007.

and yields meaningless parameter estimates[3]. In any case, this particular strategy amounts to assuming $\partial y/\partial w_t = 0$, which seems unduly restrictive when we are attempting to identify the impact on GDP of shocks to the wage curve.

Thus we are left with three identification strategies, which we label S1, S2, and S3. S1 adds $a_{12} = 0$ to (15), S2 adds $a_{21} = 0$ to (15), and S3 adds $a_{23} = 0$ to (15). These amount to assuming $\partial y/\partial L_t = 0$, $\partial l/\partial Y_t = 0$, and $\partial l/\partial w_t = 0$, respectively. The first means that any shock to labour demand does not indirectly affect aggregate demand, which is perhaps slightly over-restrictive in that increases in labour income resulting from increased employment, *ceteris paribus*, cannot affect aggregate demand within the period. The second means that any shock to aggregate demand does not indirectly affect labour demand within the period, which again is rather over-restrictive given standard production functions. The last identification strategy means that any shock to the wage curve does not indirectly affect labour demand within the period. This is, again, relatively restrictive as it assumes that the elasticity of substitution between labour and other inputs is exactly zero in the short run. In the absence of compelling evidence, we will take an ambivalent approach and consider the behaviour of the model under all three identification strategies. Note, finally, that an identification strategy in the spirit of Balmaseda et al (2000), or the New Keynesian approach more generally, would be to assume $a_{13} = 0$ and freely estimate b_{13}. This illustrates the issue of observational equivalence referred to above, although it is not in fact the approach taken in Balmaseda et al (2000), which utilises long run restrictions.

[3]The estimates for this identification strategy are considered in appendix A.

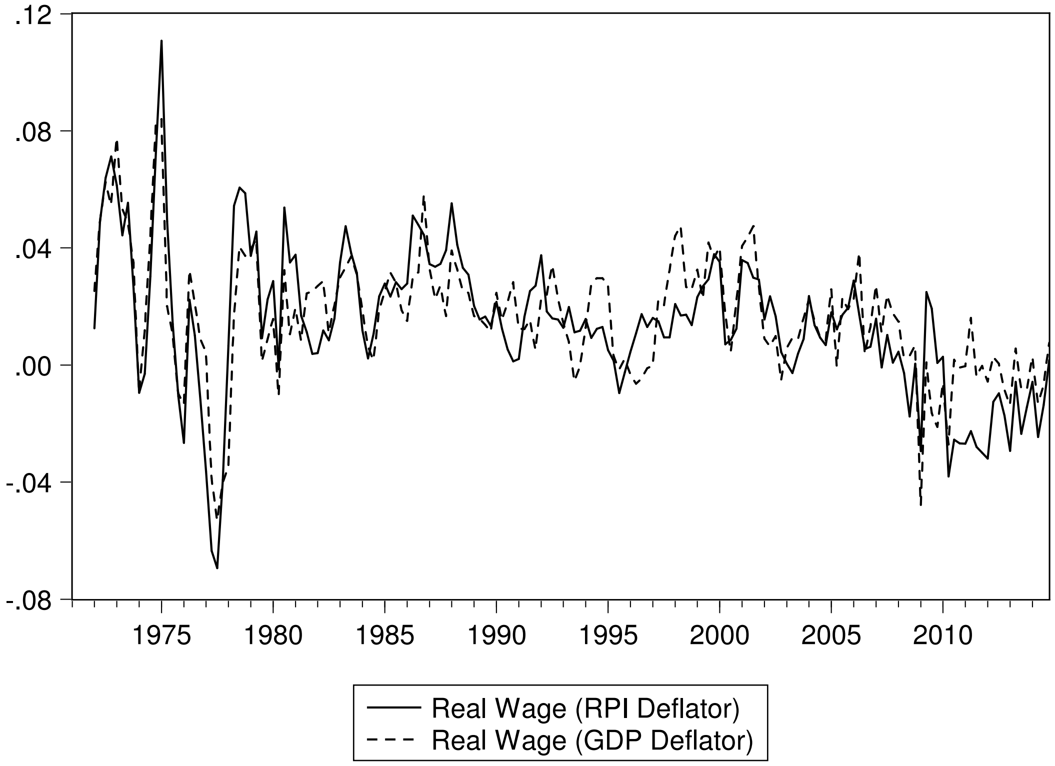

Figure 3: Alternative real wage series, 1972 - 2014.

4 Data

As explained in sections 2 and 3, we estimate SVARs in GDP, total employment, and the real wage. The raw data are as follows: quarterly nominal GDP from the UK Economic Accounts (code BKTL), UK real (chain value) GDP from the UK Economic Accounts (code BKVT), total employment from the Labour Force Survey, average weekly earnings from the ONS Employment and Earnings publications (codes MD9M and KA46), and the retail price index from the ONS MM23 Consumer Price Indices publication (code CDKO). To avoid the use of seasonal adjustment procedures as much as possible, only the total employment data is seasonally adjusted. This is because LFS data is annual up to the early 1990s, and all quarterly data prior to 1992 is interpolated between annual data on a seasonally adjusted basis. As we avoid seasonal adjustment in the remaining variables, all series are converted to year-on-year (four quarter) log differences.

Whilst the real GDP series is taken as given, we construct two alternative real wage series using average earnings deflated by the retail price index and average earnings deflated by the GDP deflator, where the latter is given by the ratio of nominal and real GDP. There is an important point to note here. Average earnings are given by the average hourly wage multiplied by average hours worked per person, which is an imperfect measure of the wage rate. Thus a positive shock to earnings could be interpreted as a positive shock to hours worked per person, which would almost certainly increase GDP. Unfortunately, the ONS average weekly earnings data does not differentiate between wage rates and hours worked, and the main source of information for this (the ASHE) is only conducted on an annual

Table 1: Linear unit root tests

	DF-GLS[a,b]	M-PP[a,b]
GDP	-2.16**	-19.11***
Employment	-1.47	-10.69**
Wage (RPI deflator)	-3.61***	-20.30***
Wage (GDP deflator)	-3.47***	-17.79***

Notes: [a]Statistics reported: DF-GLS=t-statistic; M-PP=MZa-statistic; [b]Critical values used: DF-GLS=Table 1 of Elliot et al. (1996); M-PP=Table 1 of Ng and Perron (2001). *, **, and *** respectively denote rejection of the null hypothesis at the 10%, 5%, and 1% confidence levels.

basis. However, we have good reason to believe that the bulk of fluctuations in average weekly earnings are accounted for by fluctuations in the hourly wage. Figure 2 plots the four quarter log difference of average weekly earnings deflated by the RPI deflator and the four quarter log difference of estimated hours worked per person from the Ohanian and Raffo (2012) database. The hours series is considerably less volatile than the real earnings series over the majority of the sample, and the contemporaneous correlation coefficient between the two series is negative and not significantly different from zero at the 5% level. Thus we expect the bulk of fluctuations in real earnings to be due to real wage fluctuations, and we do not expect positive shocks to real earnings to be the result of increases in hours worked.

Given this, we will treat the average earnings series as an accurate measure of average wages. At the same time, it is important to use both the RPI and GDP deflators in constructing real wages, because whilst consumer prices are of chief importance to households in wage bargaining, output prices are of chief importance to firms in wage bargaining, and thus both deflators should be used to test for robustness. In fact, at business cycle frequencies the two series diverge significantly in certain periods. Figure 3 plots the two alternative real wage series over the period 1972 - 2014, with the data in four quarter log differences. Whilst the two series appear to move together in the 1970s, most of the 1980s, and the 2000s, they diverge somewhat in the 1990s. The most significant divergence, however, is in the 2010s. Whilst both real wage series are negative in the immediate aftermath of the 2008 recession, the RPI deflated series is then negative from 2010q2 to 2014q3, whilst the average of the GDP deflated series is approximately zero over the same period.

Figure 3 highlights a more important point, however, which is the apparent mean shift in both the real wage series after 2008. Unfortunately, as this shift happens near the end of our sample, it is difficult to test for formally, and as such we will assume it exists and restrict our estimation sample to the period 1971q1 - 2007q4. Figure 4 plots four quarter log differences of real GDP, total employment, and the real wage over this estimation period, with the RPI deflated real wage series in the top panel, and the GDP deflated real wage series in the bottom panel. Two different unit root tests were employed to determine the order of integration of the series: the Dickey–Fuller Generalized Least Squares test (DF-GLS; Elliot et al. 1996), and the Modified Phillips-Perron test (M-PP; Ng and Perron

2001), using OLS-detrended data as the autoregressive spectral estimation method[4]. The highest lag order (l_{max}) selected to carry out both tests was determined from the sample size according to the method proposed by Schwert (1989): $l_{max} = \left[12\,(144/100)^{0.25}\right] \approx 13$. The optimal lag order was selected according to the Modified Akaike Information Criterion (MAIC) proposed by Ng and Perron (2001), which reduces size distortions substantially. Table 1 reports the different linear unit root tests that best capture the actual behaviour of the series in order to avoid misspecification[5]. The latter shows that the DF-GLS and the M-PP tests reject the null of a unit root in the majority of cases, the only exception being the total employment series when the DF-GLS test was employed. Hence, we can conclude that all series can be characterised as $I(0)$ processes.

Whilst the real wage appears to be mildly pro-cyclical in figure 4, and less volatile than GDP, total employment follows GDP with a considerable lag and is much less volatile than GDP, even allowing for the interpolated data prior to 1992. Given these observations, figure 5 plots four quarter log differences of real GDP and the labour share, where the latter is approximated by the four quarter log difference of the real wage plus the four quarter log difference of total employment minus the four quarter log difference of real GDP. As before, the top panel displays the labour share using the RPI deflator, and the bottom panel displays the labour share using the GDP deflator. This figure illustrates the well known counter-cyclicality of the labour share, in contrast with the mild pro-cyclicality of the real wage illustrated in figure 4. Note that the wage led aggregate demand hypothesis as stated in section 2 proposes that a positive shock to the real wage results in an increase in both GDP and the labour share, which might appear inconsistent with the counter-cyclicality of the labour share at first glance. However, all this implies is that shocks to the real wage that increase both the labour share and GDP cannot be a significant source of fluctuations in either variable, and the wage led aggregate demand hypothesis is thus perfectly consistent with a counter-cyclical labour share. This conjecture, in fact, is confirmed by the variance decomposition analyses presented in section 5.

5 Results

As discussed in sections 2, 3 and 4, we estimate SVAR models in GDP, total employment, and the real wage. We first estimate the reduced form (13) for the case in which the wage is deflated by the RPI deflator, and the reduced form (13) for the case in which the wage is deflated by the GDP deflator, using ordinary least squares. The AIC lag length criteria for both the RPI deflator model and the GDP deflator model indicate 10 lags, and 10 lags also results in reasonable statistics for residual autocorrelation. Particularly, we cannot reject the null of no residual autocorrelation at the 1% level at lags 1-12 for the RPI deflator model, and we cannot reject the null of no residual correlation at the 5% level at lags 1-12 for the GDP deflator model. Both models with 10 lags are stable, with all roots lying within the unit circle. In addition, we cannot reject the null of homoskedasticity using the White Test for either model.

Although the simple reduced form models appear to be well specified in terms of residual autocorrelation and heteroskedasticity, residual normality test results are unsatisfactory. On

[4]According to Perron and Qu (2007), this method can be considered as a solution to the drawback that, for non-local alternatives, the power of the M-PP tests can be very small.

[5]Both tests were carried out including a constant as exogenous regressors for the different series. Different specifications did not change the main conclusions.

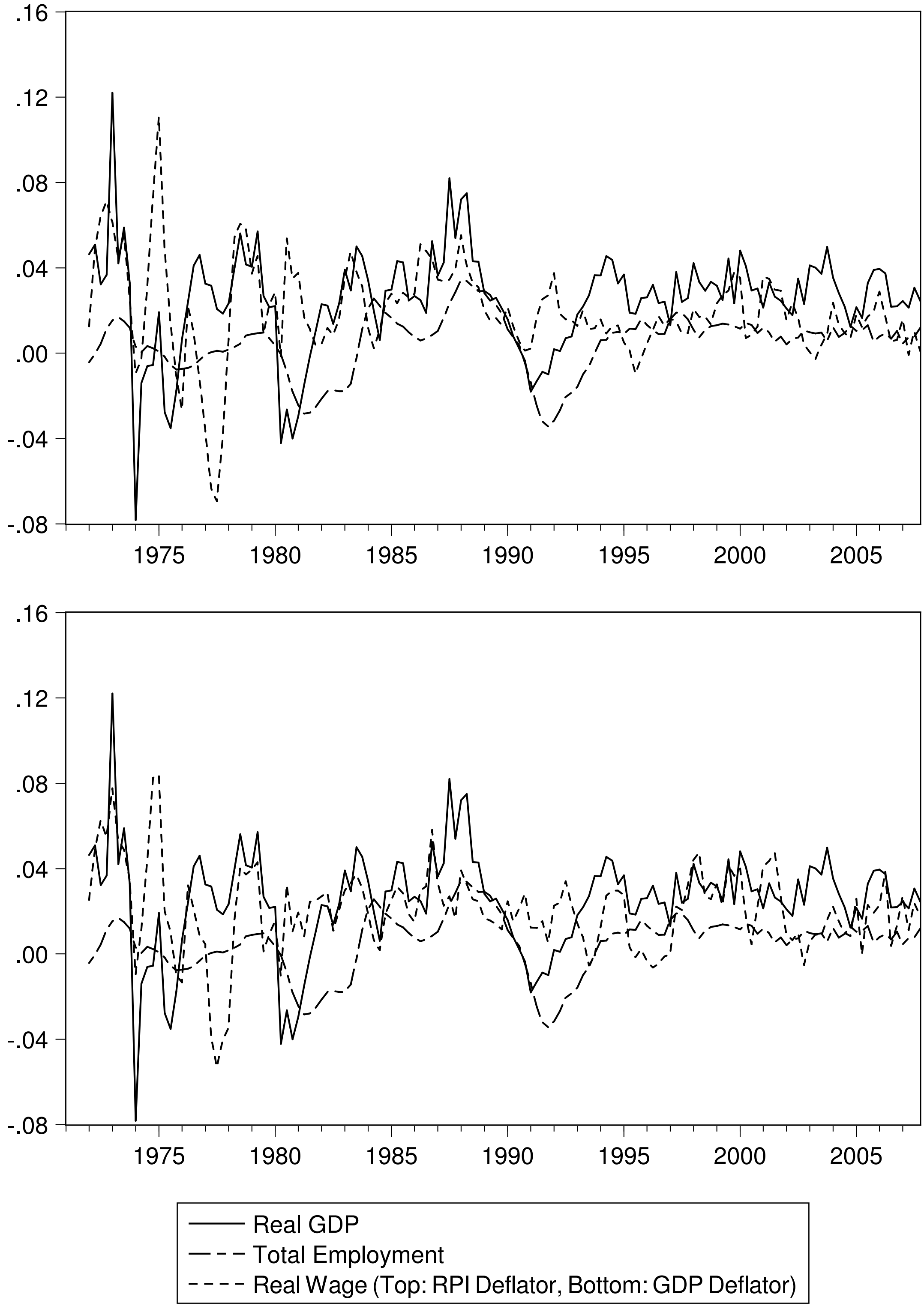

Figure 4: GDP, employment, and real wage series, 1972 - 2007.

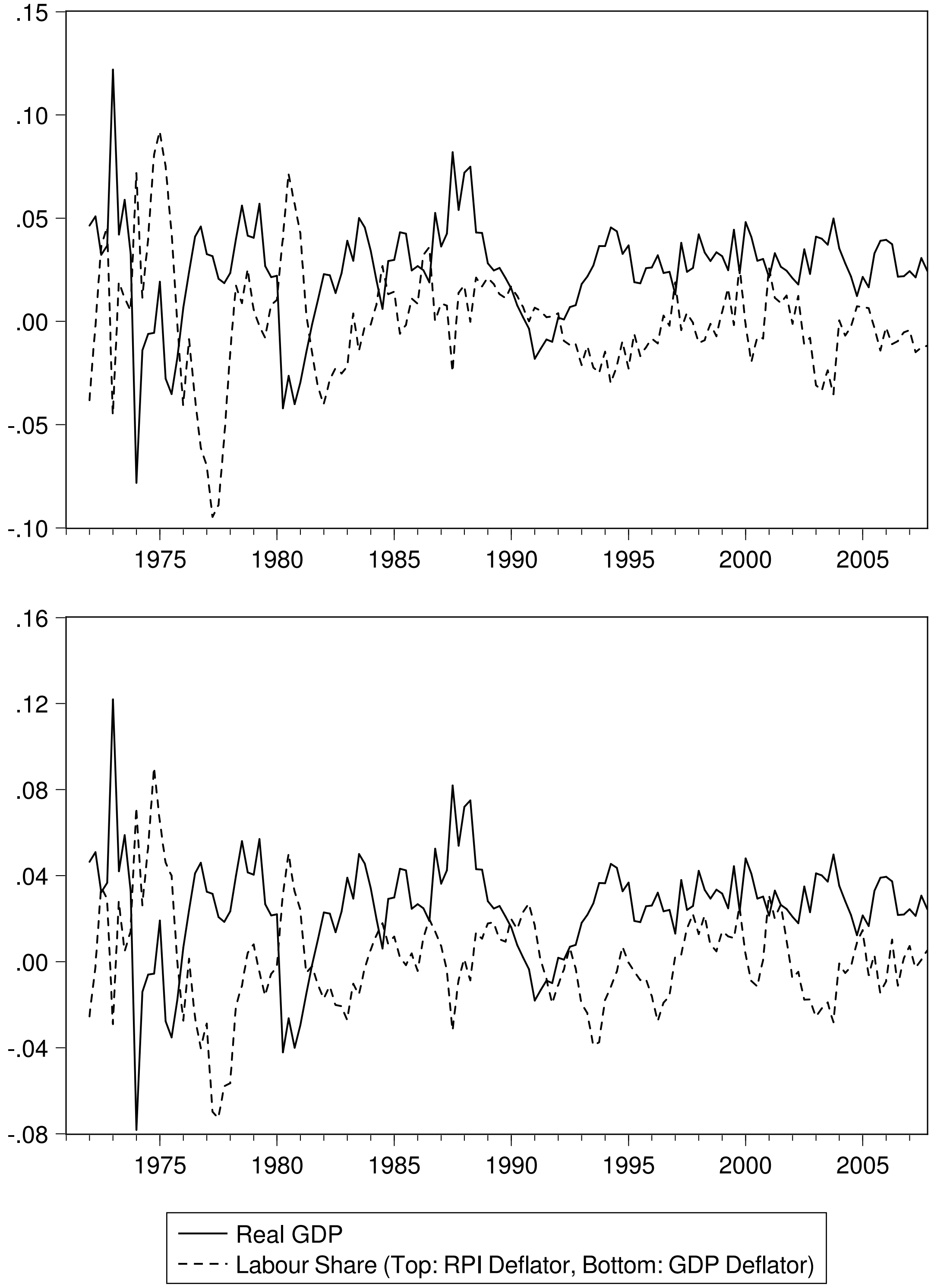

Figure 5: GDP and labour share approximation, 1972 - 2007.

inspection of the reduced form residuals, an outlier in the GDP series at 1980 quarter 2 is apparent, corresponding to the start of the recession in the early 1980s. As a result, we re-estimate the reduced form models with a dummy variable equal to 1 at 1980 quarter 2, and zero otherwise. This mitigates the residual normality problems without materially affecting the residual autocorrelation or heteroskedasticity test results. Given this, we choose 10 lags for both reduced form VARs, and proceed to estimate the structural factorisations. Detailed tables of results for the specification tests (after the inclusion of the aforementioned dummy variable) are provided in appendix B.

As discussed in section 3, we rely on three identification strategies, labelled strategies S1, S2, and S3. We estimate the A and B matrices using FIML in EViews. Our results for the model using RPI deflated wages are presented in (i) - (iii). Estimates that are significantly different from zero at the 5% level are given in boldface:

$$\text{S1:} \quad \begin{pmatrix} 1 & 0 & \mathbf{-0.24} \\ \mathbf{-0.07} & 1 & -0.00 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} \mathbf{0.009} & 0 & 0 \\ 0 & \mathbf{0.003} & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}, \tag{i}$$

$$\text{S2:} \quad \begin{pmatrix} 1 & \mathbf{-0.83} & \mathbf{-0.23} \\ 0 & 1 & -0.02 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} \mathbf{0.009} & 0 & 0 \\ 0 & \mathbf{0.003} & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}, \tag{ii}$$

$$\text{S3:} \quad \begin{pmatrix} 1 & 0.05 & \mathbf{-0.24} \\ -0.07 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} \mathbf{0.009} & 0 & 0 \\ 0 & \mathbf{0.003} & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}. \tag{iii}$$

Our results for the model using GDP deflated wages are presented in (iv) - (vi). Estimates that are significantly different from zero at the 5% level are again given in boldface:

$$\text{S1:} \quad \begin{pmatrix} 1 & 0 & \mathbf{-0.51} \\ \mathbf{-0.06} & 1 & -0.02 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} \mathbf{0.008} & 0 & 0 \\ 0 & \mathbf{0.003} & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}, \tag{iv}$$

$$\text{S2:} \quad \begin{pmatrix} 1 & \mathbf{-0.56} & \mathbf{-0.48} \\ 0 & 1 & \mathbf{-0.05} \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} \mathbf{0.008} & 0 & 0 \\ 0 & \mathbf{0.003} & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}, \tag{v}$$

$$\text{S3:} \quad \begin{pmatrix} 1 & 0.32 & \mathbf{-0.52} \\ \mathbf{-0.09} & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} \mathbf{0.008} & 0 & 0 \\ 0 & \mathbf{0.003} & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}. \tag{vi}$$

These estimates provide robust evidence that positive shocks to the real wage increase aggregate demand in the United Kingdom. For each identification strategy, the model using RPI deflated wages predicts that a 1 percentage point increase in real wage growth, as a result of a shock to the wage curve, will lead to an increase in GDP growth of approximately 0.24 percentage points within the quarter. The model using GDP deflated wages predicts a greater increase in GDP - for each identification strategy in this model, a 1 percentage point increase in real wage growth, as a result of a shock to the wage curve, will lead to an increase in GDP growth of approximately 0.5 percentage points within the quarter.

The six estimates of the B matrix are very similar to one another, whilst the only estimates that differ greatly across identification strategies and the choice of deflator are the partial effects of GDP and total employment. However, the only estimates of a_{12} and a_{21} with unexpected sign are not significantly different from zero at conventional confidence levels. It is also interesting to note that total employment does not appear to respond strongly to real wage shocks within the period, with only one estimate of a_{23} being significantly different from zero at the 5% level. Finally, the results presented in (i) - (vi) do not qualitatively change if we omit the dummy variable at 1980 quarter 2. In particular, all estimates are the same sign, with the only material differences being increases in the estimates for a_{13}.

As noted in section 2, the conclusion that positive shocks to the real wage increase aggregate demand does not, by itself, lead to the conclusion that the wage led aggregate demand hypothesis is supported. However, as total employment does not appear to respond to real wage shocks within the quarter in our models, and the response of the real wage to a wage curve shock is uniformly *greater* than the response of aggregate demand, then we can say with confidence that both aggregate demand *and* the labour share increase in response to a real wage shock in the United Kingdom. As a result, our estimates provide robust evidence in favour of the wage led aggregate demand hypothesis in the United Kingdom.

These conclusions are reinforced by the impulse response functions presented in figure 6. These plot the analytical impulse response functions and 95% confidence bands for all three variables, in response to a positive shock to the wage curve. As the impulse response functions for real wage shocks are very similar for each identification strategy, we only present the results for strategy S1[6]. Figure 6 plots the responses for the model with RPI deflated wages in the left three panels, and responses for the model with GDP deflated wages in the right three panels. The differences between the two sets of impulse response functions are minor, with the obvious difference being the larger effect of the real wage shock on GDP in the model with GDP deflated wages. Aside from this, both sets of impulse response functions imply that the positive effect on GDP growth of a real wage shock declines to zero after a year, and that the effect on total employment is negligible.

Finally, it was noted in section 4 that the labour share is counter-cyclical, despite the result that identified shocks to the real wage increase both GDP and the labour share. This implies that shocks to the real wage cannot account for a significant proportion of GDP movements, which is confirmed by the forecast error variance decompositions in table 2. From this table we arrive at three conclusions: *a.* the results obtained from the different identification strategies are fairly similar, *b.* higher proportions of the variation in GDP can be explained by real wage shocks when using the GDP deflator, and *c.* shocks to the real wage explain at most 35% of the variation in GDP (which corresponds to the first lag of the forecast horizon when the SVAR is estimated using GDP deflated wages).

[6] In fact, the impulse response functions for real wage shocks are almost identical for the different identification strategies. The impulse response functions for the remaining shocks differ somewhat between identification strategies, as implied by the differences in a_{12} and a_{21} estimates in (i) - (vi).

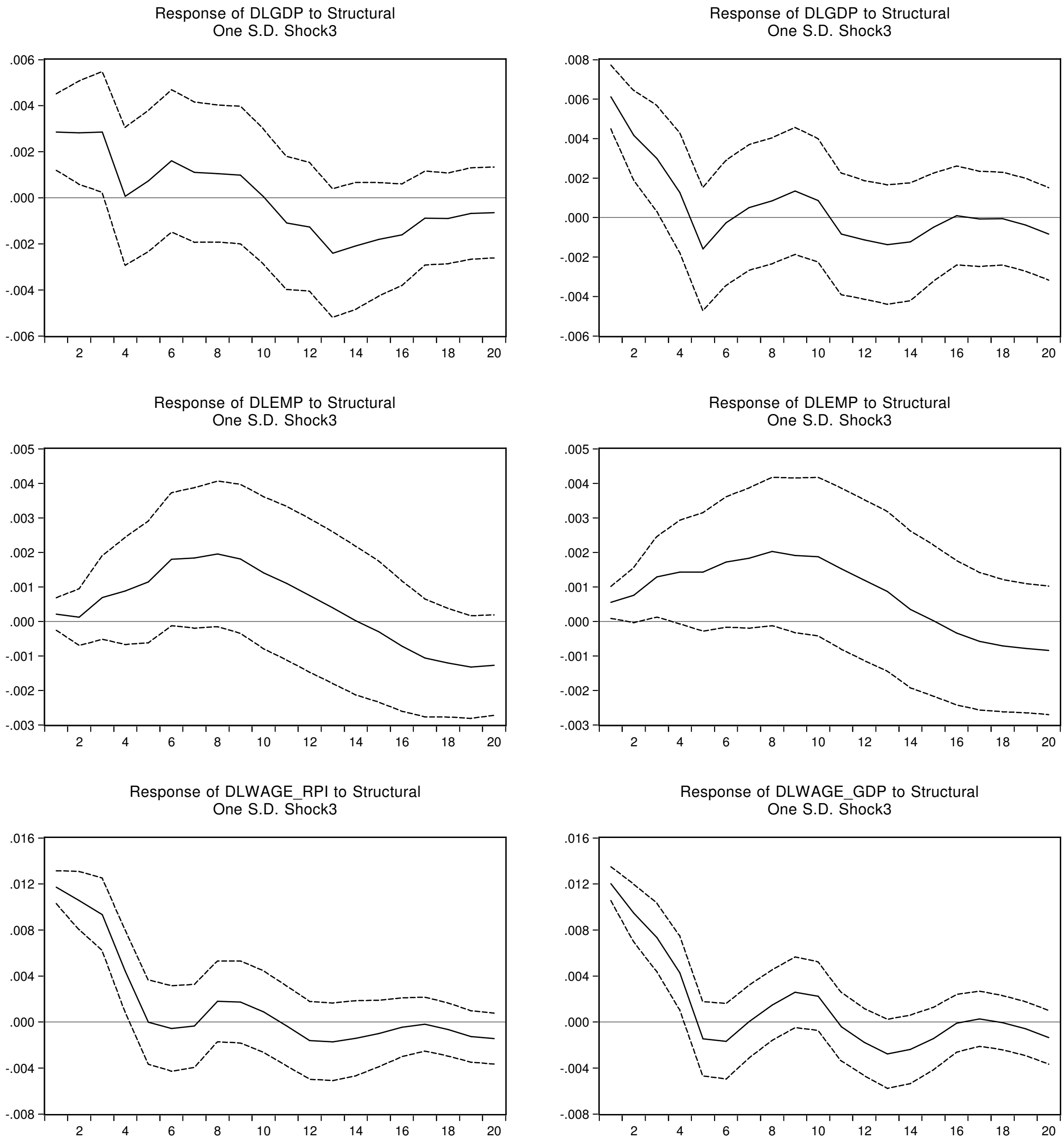

Figure 6: Impulse response functions to real wage shocks, identification strategy S1. Left panel: RPI deflated wage, right panel: GDP deflated wage.

Table 2: Structural variance decompositions of GDP[a]

Forecast Horizon	S1	S2	S3
SVAR models using RPI deflated wage[b]			
1	8.45	8.45	8.45
2	10.08	10.08	10.08
4	9.49	9.49	9.49
8	9.95	9.95	9.95
SVAR models using GDP deflated wage[b]			
1	35.29	35.29	35.29
2	31.86	31.86	31.86
4	23.36	23.36	23.36
8	20.49	20.49	20.49

Notes:[a]We only show the results of GDP to shocks in the real wage; [b]Percentage points are shown.

These results corroborate our previous conjecture that shocks to the real wage explain a relatively small proportion of the variance in GDP. Therefore, whilst positive real wage shocks result in an increase in both GDP and the labour share, these shocks do not account for a large proportion of GDP movements, such that the observed labour share in the UK is countercyclical.

6 Concluding Remarks

This paper starts with the observation that the wage led aggregate demand hypothesis has been tested in UK using a variety of estimation procedures and data types. Studies relying on equation-by-equation estimation procedures with annual data tend to find support for wage led aggregate demand in the UK, whilst studies using systems estimation procedures or quarterly data tend to find less or no support for the hypothesis. In order to resolve this incongruity, we test the wage led aggregate demand hypothesis using SVAR models estimated on quarterly data. We use a variety of data types and filtering methods, and take a relatively liberal approach to identification. Our conclusions are unambiguous: all of our models indicate that positive shocks to the real wage increase both GDP and the labour share, indicating that aggregate demand in the UK is wage led. Thus our investigation provides robust evidence in favour of the wage led aggregate demand hypothesis in the UK.

Whilst these results are encouraging, the effects of movements in the income distribution on aggregate demand remain contested. This is the case both within and without heterodox economics, with an American school of structuralist economists remaining sceptical of the hypothesis (Barbosa-Filho 2015), and New Keynesian economists largely ignoring the possibility that changes in the income distribution can have short run effects on the level of demand. Thus two main avenues for future research immediately suggest themselves:

first, applying the method of this paper to other countries in which the existing literature is inconclusive, and second, testing the wage led aggregate demand hypothesis in a New Keynesian setting. The latter requires a more sophisticated treatment of nominal wage and price inflation than we provide in the present paper, which is the area we are currently investigating.

Finally, an avenue that has not been explored in the wage led aggregate demand literature is the effect of wage curve shocks in the presence of nonlinearities (although see Nikiforos and Foley (2012) in this respect). One might conjecture that income distribution changes have asymmetric effects over the business cycle. In particular, one might imagine that redistributing income from profit earners to wage earners, and thus boosting consumption, might have a greater impact on GDP in a recession than in an expansion. Exploring this avenue using nonlinear time series models is the second area we plan to explore in the immediate future, which will hopefully shed more light on the interplay between the income distribution and aggregate demand.

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Appendix A

As discussed in section 3, we have a fourth potential identification strategy defined by equation (A1):

$$\begin{pmatrix} 1 & a_{12} & 0 \\ a_{21} & 1 & a_{23} \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} b_{11} & 0 & 0 \\ 0 & b_{22} & 0 \\ 0 & 0 & b_{33} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}. \tag{A1}$$

As in the identification strategies S1, S2, and S3, discussed in the main body of the paper, the above strategy assumes that the real wage does not react to aggregate demand or labour demand shocks within the period, and adds the restriction $\partial y/\partial w_t = 0$. This amounts to assuming that GDP does not respond to wage curve shocks within the period. This model appears to have a very flat likelihood function (along at least one dimension), and the standard EViews maximum likelihood algorithm fails to converge for most starting values. For certain values the algorithm will converge, and returns the following estimates of A and B (for the model with RPI deflated wages):

$$\begin{pmatrix} 1 & -14.2 & 0 \\ 31.3 & 1 & -11.1 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \epsilon_Y \\ \epsilon_L \\ \epsilon_w \end{pmatrix} = \begin{pmatrix} 0.037 & 0 & 0 \\ 0 & 0.326 & 0 \\ 0 & 0 & \mathbf{0.012} \end{pmatrix} \begin{pmatrix} u_Y \\ u_L \\ u_w \end{pmatrix}. \tag{A2}$$

These estimates are economically meaningless, recalling that we are dealing with log differenced data. In addition, the standard errors are extremely large, and only the estimate of b_{33} is statistically different from zero at the 5% level (and the same as its estimates in S1, S2, and S3, as this parameter is identified purely from ϵ_w). Finally, we tested a number of different estimation procedures for this identification strategy, with the most successful being the Newton-Raphson algorithm (numerical score, applied in Gretl). This converged reliably, with parameter estimates very similar to those given in equation (A2).

It is possible that both the EViews scoring algorithm and the Gretl Newton-Raphson algorithm are both locating the same *incorrect* local maximum. However, it is also possible that this model is just badly identified, and is producing correspondingly meaningless results. As noted in section 3, above, we proceed on the basis of the latter, given that $\partial y/\partial w_t = 0$ is an odd assumption given the purpose of our investigation. As noted in the conclusion, our research continues on this topic, and may shed light on this particular problem in the future.

Appendix B

We estimate two different VAR models: one using RPI deflated wages and the other one using GDP deflated wages. Both VAR models are stable (their roots have modulus less than 1 and lie inside the unit circle), and we report the main misspecification tests in Table 3:

Table 3: Joint misspecification tests over the VAR models[a]

Autocorrelation[b]		Heteroskedasticity		Normality	
Statistic	*p*-value	χ^2 statistic	*p*-value	Statistic	*p*-value
VAR using RPI deflated wages					
23.88	0.01	377.11	0.33	4.06	0.67
VAR using GDP deflated wages					
14.38	0.11	281.52	0.99	19.46	0

Notes: [a]Tests employed: Serial correlation=Lagrange Multiplier; Heteroskedasticity=White (no cross terms); Normality=Cholesky of covariance (Lutkepohl); [b]We only report the results that test for first-order serial correlation. The Lagrange Multiplier statistics of the second-order serial correlation tests are the following: VAR using $w_{1,t}$=7.53 (p-value=0.58); and VAR using $w_{2,t}$=10.24 (p-value=0.33).

From Table 2 it is possible to observe that neither VAR model presents problems of autocorrelation, heteroskedasticity or normality at the 1% level of significance, the only exception being the VAR model using $w_{2,t}$ (which presents problems of normality).