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Labour Market Effects of Eastern European Migration in Wales

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The enlargement of the European Union in May 2004 triggered a relatively large and rapid migration inflow into Wales which was concentrated into narrow areas and occupations. As this inflow was larger and faster than anticipated, it arguably corresponds more closely to an exogenous supply shock than most migration shocks studied in the literature. This helps to some extent to circumvent identification issues arising from simultaneity bias which usually pose difficulties when estimating the effect of migration inflows on the labour market. We found little evidence that the inflow of accession migrants contributed to a fall in wages or a rise in claimant unemployment in Wales between 2004 and 2006. In particular, we found no evidence of an adverse impact on young, female or low-skilled claimant unemployment and no evidence of an adverse impact on the wages of the low-paid. If anything, we found a positive effect on the wages of higher paid workers and some weak evidence of a potentially favourable impact on claimant unemployment.

JEL classification: J22.

Keywords: migration, employment, wages, Central and Eastern Europe, UK, Wales.

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1. Introduction

The enlargement of the European Union (EU) in May 2004 granted workers from ten accession countries free movement within the union. This triggered a relatively large and rapid migration inflow into Wales which was concentrated into narrow areas and occupations. As Wales has historically experienced both net out-migration of younger people and in-migration of older people, inflows of younger workers has been seen as one solution to problems caused by an ageing workforce (Drinkwater and Blackaby 2004). Thus, the analysis of the effects of the recent accession migration inflow on the Welsh labour market provides a particularly interesting case study.

Around 16,000 accession migrants joined the Welsh labour market between May 2004 and May 2006, according to the Worker Registration Scheme (WRS). This migration inflow was equivalent to roughly 1% of the Welsh working age population and to roughly a third of the Jobseeker's Allowance (JSA) claimant unemployment by May 2006. Therefore, this migration shock was large enough to have adverse effects on the labour market. Coincidently, claimant unemployment rose by roughly 5,000 (or 12%) between May 2004 and May 2006.

The main contribution of this paper is to investigate whether this increase in the claimant unemployment was due to the accession migration inflow. More specifically, we estimate the effect of the accession inflow on claimant unemployment and on the distribution of wages using micro level monthly WRS and JSA data, as well as data from the Annual Survey on Hours and Earnings (ASHE) between 2004 and 2006. This new and large source of data on migration combined with data on claimant unemployment permits disaggregation at fine (district and month) levels and offers invaluable insights into the Welsh labour market.

Given that the paucity of suitable data is one of the main reasons for scarce evidence on the effects of migration, this paper helps to fill a gap in the migration literature for Wales and the UK, which is very limited (Dustmann et al. 2005 and 2007; Manacorda et al. 2006; Lemos and Portes 2008; Drinkwater et al. 2009 are some of the few) – specially so on the effects of the recent EU enlargement. Furthermore, this paper helps to inform policymaking on the face of further EU enlargement. It is especially opportune, given the current heated public debate on migration – and in particular on migration from current and future accession countries.

Another contribution of this paper is that the nature of the accession migration helps to some extent to circumvent identification issues arising from simultaneity bias which usually pose difficulties when estimating the effect of migration inflows on the labour market. One complicating identification issue is that if natives respond to the migration inflow by moving away from a particular area or occupation, then potential adverse effects on that labour market may be offset. Another complicating identification issue is that if migrants respond to specific demand conditions by self-selecting into particularly booming areas or occupations, once again potential adverse effects on that labour market may be offset. The nature of the accession migration, however, was such that these responses from both natives and migrants might have been sufficiently lagged to allow identification of adverse wages and unemployment effects. That is because the accession inflow was substantially larger and faster than anticipated (see Dustmann et al. 2003 for forecasts), and thus arguably corresponds more closely to an exogenous supply shock than most migration shocks studied in the literature.

We found little evidence that the inflow of accession migrants contributed to a fall in wages or a rise in claimant unemployment in Wales between 2004 and 2006. In particular, we found no evidence of an adverse impact on young, female or low-skilled claimant unemployment and no evidence of an adverse impact on the wages of the low-paid. If anything, we found a positive effect on the wages of higher paid workers and some weak evidence of a potentially favourable impact on claimant unemployment. Our results are robust to a number of specification checks and are in line with other results in the literature.[1]

Our results are also in line with standard theory, which predicts adverse wages and/or employment effects following a migration inflow that is unbalanced across areas or skills. We found evidence that higher paid (complement) workers had larger, positive and significant wage increases, whereas lower paid (competing) workers had smaller and insignificant wage increases. One interpretation here is that, relative to higher paid workers, lower paid workers had less favourable (though not adverse) wage increases. Incidentally, more adverse wage effects for lower paid (competing) workers may have been potentially mitigated or offset because they were protected by a concurrently increasing minimum wage.

[1] Our results are in line with evidence in the international (mainly US) literature of little or no effect on employment and wages (Chiswick 1980; Grossman 1982; Card 1990, 2005 and 2007; Altonji and Card 1991; Pischke and Velling 1997; Friedberg 2001; Dustmann et al. 2005 and 2007; Manacorda et al. 2006; Carrasco et al. 2008), though in contrast with other evidence of more adverse effects (Borjas 2003 and 2006; Angrist and Kugler 2003; Orrenius and Zavodny 2007). As we discuss in Sections 3 and 4, the disagreement in the literature is underlined by an ongoing debate on identification issues arising from natives' mobility and migrants' self-selection (see for example Chiswick 1991; Borjas 1999; Card 2001).

The remainder of this paper is organised as follows. In Section 2 we describe our data. In Section 3 we discuss our empirical model of unemployment. In Section 4 we carefully discuss our empirical approach and several associated identification issues. In Section 5 we discuss our empirical model of wages. In Section 6 we summarise and conclude.

2. Data

2.1 Sources

The migration data we use is from the WRS, the unemployment data is from the JSA, and the wages data is from the ASHE. We discuss each in turn.

In order to work in the UK for a month or longer, accession nationals are obliged to register on the Home Office administered Worker Registration Scheme (WRS). Registration, in addition to being a legal requirement for accession migrants, offers incentives such as certain social security benefits (Home Office 2004). As a result, compliance is high and we observe all those registered on the WRS. Between May 2004 and May 2006 [May 2004 and December 2008], around 560,000 [870,000] migrants registered, according to the WRS [UK Border Agency]. In Wales, these figures were 16,000 and 25,000 respectively. The left panel of Figure 1 shows the quarterly inflow between May 2004 and December 2008 according to the UK Border Agency.[2] The trend shows a seasonal pattern where numbers peak in the summer and plunge in the winter. This trend is also observed in the WRS data (see Figure 3). Furthermore, a downwards trend can be observed from 2008 onwards.[3] The migrant headcount is relatively small for Wales when compared with other parts of the UK (also see Figure 3). This is also illustrated in Figure 2, which shows that Wales received a smaller migration inflow relative to its working age population than other parts of the UK.

The WRS data is rich, large, frequent and timely. It records nationality, address, age,

[2] The Home Office (2006) uses "application date" and the UK Border Agency uses application "approval date" to aggregate the data, whereas Gilpin et al. (2006) use "entry date". As the typical migrant enters the UK, finds a job, and then applies to the WRS, we use "start of work date" to best capture labour market effects and to skew from identification issues associated to using "entry date" or "application date".

[3] In our regression analysis in Sections 3 to 5 we use WRS data from May 2004 to May 2006, whereas in some of our descriptive analysis in Section 2 we also use UK Border Agency data from May 2004 to December 2008. This is primarily because, although our first request for WRS and JSA monthly micro level data was successful in 2007, our second request in 2009 was not. Therefore, to gather a sense of more recent descriptive figures we use quarterly UK Border Agency data, which displays lower overall numbers. In addition to exploiting the better quality of the WRS data, further arguments for restricting the regression analysis to the first two years only are that: in the longer run labour markets adjust (which might dilute potentially adverse effects) (Altonji and Card 1991; Dustmann et al. 2005); and in the longer run labour markets are hit by other shocks (which makes identification of migration effects more challenging).

gender, number of dependents, application date, entry date, start of work date, hourly wage rate, hours worked, sector, occupation and industry. Table 1 shows that many WRS migrants are young, male, Polish, childless, working full time in low-paid jobs in elementary and machine operative occupations and in manufacturing and catering. The WRS is only available for migrants from the ten accession countries, as migrants from other countries are not required to register. We restrict our sample to eight of those (A8), namely: Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovenia and Slovakia. We exclude Malta and Cyprus, which already had relative access to the EU labour market.[4]

In order to claim unemployment benefit, workers are obliged to register on the Department for Work and Pensions administered Jobseeker's Allowance (JSA) programme. Registration is a legal requirement to qualify for the benefit, and therefore compliance is full and we observe all those receiving JSA. Between May 2004 and May 2006 [May 2004 and December 2008], JSA claimant unemployment rose by roughly 96,000 [690,000]. In Wales, these figures were 5,000 and 40,000 respectively. Figure 3 shows the monthly JSA stock during this period. As before, the trend shows a seasonal pattern during the summer and winter. Claimant unemployment was roughly stable in 2005; it increased in 2006, and then decreased in 2007, despite a continuous inflow of migrants. It then increased sharply from 2008 onwards, which coincides with the current financial crisis.

The JSA data is large, frequent and timely, and like the WRS, permits disaggregation at fine (district and month) levels.[5] This is in contrast with the more widely used Labour Force Survey (LFS), where migration analysis below the region and quarter level is not feasible due to sample size limitations. Furthermore, the JSA measures claimant unemployment, which is directly relevant for policymaking, instead of the more broad ILO unemployment. The JSA records address, gender, age, usual and sought occupations, claim

[4] A caveat with the WRS is that it measures inflows only and thus the associated netflow and stock cannot be calculated. That is because the WRS records jobs, not people. Migrants leaving are not counted whereas migrants re-entering the UK are double counted. Blanchflower et al. (2007) analyse A8 migration figures across several data sources and conclude that a stock of 500,000 migrants by late 2006 is likely to be an upper bound (for example, among other reasons, the WRS data shows a high proportion of registrations for temporary jobs and few migrants indicating "likely length of stay" over an year). Pollard et al. (2008) and Coats (2008) provide similar analysis and conclude that outflow is not zero, in line with evidence on return migration (LaLonde and Topel 1997). If outflow is not random, β^n in Equation 1 could be biased (see Section 2.2). Gilpin et al. (2006) provide a detailed discussion on measurement error in the WRS and conclude that any associated bias is not too severe. Another caveat with the WRS is that registration is not a requirement for the self-employed (who are a minority that already had relative access to the EU labour market), which explains the larger number of Polish plumbers in anecdotal evidence.

[5] We use ONS-defined geographical areas: 409 Local Authority Districts, 49 counties and 12 Government Regions (ONS 2003) (see Table 1). Wales comprises 22 Unitary Authority Districts (also see Section 4.2).

start and end dates. It does not, however, record nationality. Table 1 shows that many JSA unemployed are over 35 years old, female and work in low-paid jobs in elementary occupation.

The Annual Survey of Hours and Earnings (ASHE), collected by the Office for National Statistics (ONS), is derived from employers' data and represents 1% of all employees, containing around 160,000 responses per tax-year (which runs from April to March). Its sample size again permits disaggregation at fine (district and year) levels, in contrast with the LFS, as discussed above. It collects, among other variables, address, gender, age, hourly pay, hours worked, occupation and industry. Table 1 shows various percentiles and the average of the ASHE and WRS hourly wage distributions. Figure 4 shows both distributions for those earning £7 per hour or below. A striking feature is how sizeable the spike at the minimum wage in the WRS distribution is in relation to the spike in the ASHE distribution, though caution should be taken here, as ASHE includes WRS migrants after 2004. Another striking feature is how remarkably compressed the WRS distribution is: over 90% (75%) of migrants earn between £2.00 (£4.00) and £7.00 an hour.

Finally, we define control variables that describe the natives' population from the LFS, collected by the Office for National Statistics (ONS). ("Natives" here and throughout the paper include UK born and overseas born nationals who are UK residents.) The LFS is a rotating panel survey that interviews around 60,000 households with about 140,000 respondents every quarter and represents 0.5% of the population. It collects information on personal characteristics and labour market variables. Table 1 summarises some variables from the LFS between April 2004 and June 2006.

2.2 Descriptive Analysis

Within Wales, WRS migrants are concentrated in five main districts, as illustrated in the right panel of Figure 1: Newport, Cardiff, Carmarthenshire, Wrexham and Flintshire. These districts form two clusters of relatively bigger cities bordering England (except for Carmarthenshire, which is in West Wales and is predominantly rural) that are historically more associated with migration and have a long term trend of out-migration (Drinkwater and Blackaby 2004). These clusters are also traditionally linked to manufacturing, though the service sector has been growing in Newport and Cardiff. The left panel of Figure 5 shows that apart from Cardiff, these are not areas of particularly high or low unemployment. (Within Cardiff, it is possible that WRS migrants shun away from high or low unemployment areas too.)

Given the disproportionate numbers of WRS migrants and claimants in Cardiff, and to a lesser extent in the other four districts, it is likely that both groups compete for the same jobs and therefore two obvious questions arise. The first question is whether migrants pushed natives out of their jobs or made it harder for them to go back into jobs in these districts. The right panel of Figure 5 shows a continuing inflow of migrants and an upwards trend on the number of claimants in these areas. This provides some evidence of an adverse association between WRS migration and claimant unemployment in Newport and Cardiff, but perhaps less evidence for Carmarthenshire, Wrexham and Flintshire.

The second question is whether migrants' inflow depressed wages. Average wages in Wales increased by 4% between May 2004 and May 2006, whereas the wages of the low-paid (at the 5th and 10th percentiles of the wage distribution) increased by 4.6% and 3.9%, and the wages of higher paid workers (at the 70th percentile), by 3.1%. This provides little evidence that the WRS migrants depressed the wages of the low-paid, relatively to high-paid, despite being disproportionately concentrated in low-skilled jobs. More generally, Table 1 shows that wages grew strongly in the bottom half of the distribution during the whole period – this wage growth was stronger in Wales than in the UK. This again provides little evidence of an adverse association between WRS migration and wages growth in Wales.

Figure 6 shows that WRS migrants are concentrated in elementary (36%) and machine operative (49%) occupations, and in the manufacturing (48%) and the distribution hotels and restaurants (23%) sectors (see Table 1).[6] Once again, given the disproportionate numbers of WRS migrants and claimants in these occupations, it is likely that both groups compete for the same jobs. The obvious question is again whether migrants pushed natives out of, or made it harder for them to go back into jobs in these occupations. The left panel of Figure 7 shows that despite the continuing inflow of migrants into machine operatives occupations, more claimants switched to this from other (usual) occupations.[7] Also, wages grew faster in machine operatives between 2005 and 2006 (3.8%) than in elementary (2.7%) or other occupations (3.5%) for the whole of the UK. This suggests that demand side factors may have driven both migrants and claimants into machine operative jobs.

The right panel of Figure 7 also shows a continuing inflow of migrants into elementary occupations, where they were probably more able to find jobs because of language or other labour market barriers (Card and DiNardo 2000; Friedberg 2001; Drinkwater et al. 2009).

[6] We use the nine Standard Occupation Codes (see Table 1).
[7] We observe both usual and sought occupation for the claimant unemployed, thus overcoming a common difficulty in the literature, where occupation is often not observed (Card 2001).

This is also the usual occupation for most claimants (34%) and Figure 7 shows that some of them switched from looking for jobs in (usual) elementary to other (sought) occupations. The switch could be because natives were pushed out of their jobs, which would suggest some evidence of an adverse association between WRS migration and claimant unemployment in elementary occupations. However, the switch could also be because of other factors, including occupational progression, sectoral or occupational shocks, macro shocks, etc., which we account for in our empirical models in Sections 3 to 5. An example of such shocks, as discussed above, is the claimant unemployment increase across all occupations in early 2006, which hints at macro effects in addition to any WRS migration effects.

The top left panel of Figure 8 plots our claimant unemployment (netflow) rate variable ΔN_{it} against our migration (inflow) rate variable ΔM_{it} across t months (May 2004 to April 2006) and i districts (i districts are replaced with j occupations in the top right panel of Figure 8). This again provides little evidence of an adverse association between the two variables. The raw data suggests that claimant unemployment did not grow faster in districts and occupations that received relatively more migrants. The two bottom panels of Figure 8 also plot the average (and 10th percentile) of the distribution of log hourly pay W_{iy} in first-difference across y tax-years (2004 to 2006) and i districts against the (April to March) yearly migration rate ΔM_{iy}. Again, this provides little evidence of an adverse association between the two variables. The raw data suggests that wages did not grow slower in districts that received relatively more migrants.[8]

In sum, the inflow of WRS migrants in Wales represents a relatively large, rapid and concentrated shock into two main occupations and five main districts, with the remainder

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[8] We define $\Delta N_{it} = \frac{\Delta N_{it}^*}{P_{it}}$ and $\Delta M_{it} = \frac{\Delta M_{it}^*}{P_{it}}$, where N_{it}^* is the number (stock) of JSA claimants, M_{it}^* is the number (stock) of WRS migrants, and P_{it} is working age population. As discussed in Section 2.1, whereas we observe the stock of claimants and can calculate the netflow of claimants as $\Delta N_{it}^* = N_{it}^* - N_{it-1}^*$, we do not observe the stock of migrants. We therefore re-define the netflow of migrants as $\Delta M_{it}^* = I_{it} - O_{it}$, where I_{it} is inflow and O_{it} is outflow of migrants. As we do not observe outflow, we again re-define $\Delta M_{it}^* = I_{it}$, as it is common in the literature (see for example Card 2001; Dustmann and Glitz 2005), and interpret it as a variable in differences. Similarly, we define natives' netflow rate as $\Delta A_{it} = \frac{\Delta A_{it}^*}{P_{it}}$ and $\Delta A_{it}^* = I_{it}^A - O_{it}^A$, where I_{it}^A is inflow and O_{it}^A is outflow of natives. We also run robustness checks where our migration and unemployment variables in Equation 1 were not standardised (re-defining $\Delta N_{it} = \Delta N_{it}^*$ and $\Delta M_{it} = \Delta M_{it}^*$) and found qualitatively similar results (also see Section 3).

occupations and districts offering clear counterfactuals. There is some weak evidence that in these main occupations and locations the inflow might be associated with adverse unemployment effects, though less evidence that it might be associated with adverse wage effects. We exploit the variation in these occupation and location choices across months to separately identify the effect of the migration shock from the effect of other supply and demand shocks on claimant unemployment and wages in our empirical models, as we discuss in Sections 3 to 5.

3. Unemployment Effects

Using a reduced form equation grounded on standard theory (see for example Borjas 1999; Card 2001; Dustmann et al. 2005), we estimate the effect of the WRS migration inflow on claimant unemployment netflow in Wales:

$$\Delta N_{it} = \beta^n \Delta M_{it} + \lambda^n \Delta X_{it} + f_t^n + \Delta\varepsilon_{it}^n \qquad (1)$$

where ΔN_{it} and ΔM_{it} are our unemployment and migration variables, defined in Section 2.2, X_{it} are labour supply and demand shifters, f_t^n is time fixed effects, and ε_{it}^n is the error term in district $i = 1,...,22$ and month-year $t = 1,...,24$. The interpretation of our coefficient of interest is that a one percentage point increase in the migration rate changes the claimant unemployment rate by β^n percentage points.

As we estimate Equation 1 in first-difference, district fixed effects were differenced out. This enables us to separate the effect of district specific factors, which might make a particular district more attractive to migrants or natives or both (such as more schools, more housing, higher wages, etc.), from the effect of the WRS shock on claimant unemployment. We model time fixed effects using 24 month-year dummies. This enables us to separate the effect of other macro shocks (such as seasonal shocks, national and international shocks, etc.) from the effect of the WRS shock on claimant unemployment.

We also control for supply and demand shifters. This enables us to separate the effect of supply and demand shocks from the effect of the WRS shock on claimant unemployment. Controls in X_{it} include the proportion of the total population who are women, young (those between 18 and 24 years of age), and ethnic minorities and migrants from outside the A8 countries. This enables us to control for higher unemployment in a particular district due to the presence of relatively more women, young, minorities or other migrants – which are groups who often experience high unemployment. Further controls include the lagged proportion of WRS migrants who are women, young and parents (along with average number of children). We also control for the lagged average hours worked by WRS migrants to account for potentially higher claimant unemployment in districts where migrants work longer hours (which may increase substitutability). We also include the lagged proportion of WRS migrants in elementary and machine operative occupations to control for occupation-district specific shocks affecting claimant unemployment. Finally, we include the lagged proportion of unemployed who are women and young, and lagged average claim duration. Lagged claim duration accounts for higher unemployment in districts with historically long spells of unemployment; it also alleviates problems arising from serial correlation in the residuals and it can be interpreted as a measure of labour demand.[9]

We perform a Generalized Least Square (GLS) correction to account for the relative importance of each district and for heteroskedasticity arising from aggregation. Also, we correct the standard errors for serial correlation across and within districts.[10] Given such stringent specifications, and given the clear counterfactuals discussed in Section 2.2, we argue that the remaining variation in the claimant unemployment rate is likely due to changes in the WRS migration inflow – and this ensures the identification of β^n.

Table 2 shows our β^n estimates. The UK results are borrowed from Lemos and Portes (2008) and are provided for comparison and completeness, but the main analysis here focuses on the results for Wales. Row 1 of Panel A shows an insignificant -0.115 (unweighted OLS) β^n estimate, which corresponds to the raw data in Figure 8. Row 2 shows an insignificant 0.024 (GLS) estimate when we control for district fixed effects. Row 3 shows an insignificant 0.006 estimate when we control for month-year fixed effects. Row 4 shows an insignificant 0.012 estimate when we control for other supply and demand shocks (which are, in the main, significant and of the expected sign here as well as in the remainder models in the paper). These estimates are numerically close to zero and statistically indifferent from zero. Thus, our results suggest little evidence of adverse

[9] As in Gilpin et al. (2006), we experimented with two types of dynamics (lagged migration rate and lagged claimant unemployment rate), which, however, did not alter our main result. Although dynamics allow for lagged adjustments due to slow responses in employment, migration effects are generally expected to be lower in the longer run than in the shorter run (Altonji and Card 1991; Dustmann et al. 2005).

[10] The appropriate weight here is the sample size used to calculate the dependent variable (working age population), but our estimates were also robust to using total population as weight instead – which reduces concerns of a potential correlation between the weight and the dependent variable affecting the results. (Also, as discussed in Section 2.2, we run robustness checks where our unemployment and migration variables were not standardised and found qualitatively similar results.) Our estimates were also robust to using, in turn, April 2004 working age population and April 2004 total population as time-invariant weight.

claimant unemployment effects at the district level.

4. Identification

4.1 Simultaneity Bias

Two main sources of endogeneity could be biasing our β^n estimates in Section 3: natives' mobility and migrants' self-selection. On the one hand, potentially more adverse effects on a particular district that received a migration inflow might be offset if natives avoid competing with migrants by moving away to other districts. On the other hand, potentially more adverse effects on a particular district that received a migration inflow might be offset if migrants deliberately self-selected into booming districts. Therefore, the extent to which any adverse unemployment effects can be identified depends on how mobile natives are across districts in response to migration inflows and on how able migrants are to self-select into booming districts.

As the WRS migration shock was substantially larger and faster than anticipated (see Dustmann et al. 2003 for forecasts), both natives' and migrants' responses – through, respectively, mobility out of and self-selection into specific districts – might have been sufficiently lagged to allow identification of adverse labour market effects. That is, the WRS inflow arguably corresponds more closely to an exogenous supply shock than most migration shocks studied in the literature (also see Card 1990 and 2007; Hunt 1992; Carrington and Lima 1996; Friedberg 2001). Because of this, we argue that any simultaneity bias is not too severe in our estimates in Section 3.

One way to check the extent of any such a bias in our estimates is to explicitly control for natives' mobility in Equation 1. This allows us to separate the effect of the WRS shock on claimant unemployment from the effect of natives moving away from (or refraining to move into) a district. In other words, we build, to some extent, a counterfactual of how mobile natives would have been in the absence of the migration inflow. Therefore, this helps to correct for both natives' mobility (omitted variable) bias and migrants' self selection (omitted variable) bias.

Ideally, we want to use a variable that measures what would have been the observed natives' net migration had migrants not arrived. As such a counterfactual is not observable, we add two observable proxies to ΔX_{it}, in turn. The first proxy is lagged working age population growth (Borjas et al. 1997; Borjas 2006) – which incidentally ensures that the variation in ΔM_{it} that identifies β^n comes from the numerator (migration inflow) and not from the denominator (working age population) (Borjas 2003). To avoid repeating the dependent variable as a regressor, we use lagged working age population growth by education group (Dustmann et al. 2005; Borjas 2006).[11] The second proxy is the native netflow rate ΔA_{it} , defined in Section 2.2.

Rows 5 and 6 of Panel A of Table 2 show that, controlling for respectively lagged working age population growth and native netflow rate does not alter the estimates qualitatively: the estimates remain insignificant and are, respectively, 0.024 and 0.017 (compare with row 4). This offers little evidence that natives' mobility offset potentially more adverse claimant unemployment effects.

4.2 Aggregation Level

Another way to check the extent of any natives' mobility bias in our estimates in Sections 3 and 4.1 is to aggregate the data at broader levels. Ideally the level of data aggregation should conform to the actual radius of job search for natives competing with migrants. However, as the boundaries of the actual radius of job search for natives are an empirical matter, we experiment with several levels of aggregation (i.e. several degrees of natives' mobility), allowing the search to take place on ever wider labour markets (Borjas 2006). We start with a Twenty-Two-Way district aggregation, followed by a Seven-Way and a Three-Way district aggregation.[12] This allows us to assess whether natives are district-bound or whether they are mobile across (nearby) districts. If natives' mobility is not exacerbated by the migration inflow, estimates at the three different levels of aggregation should not differ much, as we now explain in detail.

In Sections 3 and 4.1 we assumed that there are 22 closed labour markets in Wales (i.e. 22 x 24 cells). While districts are unlikely to exactly coincide with local labour markets, they may represent a fairly realistic practical radius of job search for the low-skilled. Because WRS migrants concentrate in low-paid jobs, they compete with low-skilled natives, who are less mobile as moving costs might be prohibitive. This effectively means

[11] We use three groups: those with a degree or above, those with GCSE or below, and those in between. The last was omitted in alternative robustness checks, which did not qualitatively alter the main results.

[12] We start with the 22 Unitary Authorities districts defined by the ONS (2003) (see Section 2.1). We then aggregate these 22 districts into 7 areas: North-West Wales (Anglesey, Conwy, Denbighshire and Gwynedd), North-East Wales (Flintshire and Wrexham), Mid-West Wales (Carmarthenshire, Ceredigion and Pembrokeshire), Mid-East Wales (Powys), South-West Wales (Bridgend, Neath Port Talbot, Swansea and Vale of Glamorgan), South Wales (Blaenau Gwent, Caerphilly, Cardiff, Merthyr Tydfil, Newport, Rhondda, Cynon, Taff and Torfaen) and South-East Wales (Monmouthshire). We finally aggregate these into 3 areas: North Wales, Mid Wales and South Wales.

that they compete in a relatively more closed market.[13]

We then allow natives further mobility by assuming that there are seven [and then three] independent and closed labour markets in Wales (i.e. 7 [3] x 24 cells). Natives can now respond to the WRS inflow by moving or commuting within seven broader labour markets, instead of being locked into 22 narrowly defined independent labour markets. The underlying assumption is that the Three-Way is realistically a more closed labour market than the Seven-Way, which is realistically a more closed labour market than the Twenty-Two-Way aggregation. Thus, if natives are relatively district-bound, then estimates at the three different levels of aggregation should not differ much. That is because the natives' mobility bias is larger the greater the degree of natives' mobility (Borjas 2006). Our final level of aggregation is the national-occupation level, as we discuss in Section 4.4, which scrapes all boundaries allowing natives mobility within a fully closed national labour market.

In contrast with the estimates in Panel A of Table 2, the estimates in Panels B and C turn negative, though they are only significant at the broadest level of aggregation in panel C. This again suggests little evidence of adverse claimant unemployment effects.

The estimates are larger the broader the aggregation level, offering, perhaps tentatively, some weak evidence that natives' mobility in response to migrants' inflow has a favourable (not adverse) effect on unemployment. Nonetheless care should be taken here, as although larger estimates might be expected at wider aggregation levels as a result of theoretical predictions regarding natives' mobility (Borjas 2003 and 2006), they might also be expected as a result of modelling choices (Peri and Sparber 2008). One example is that three area dummies do not control for as many area specific shocks as 22 area dummies do, which may result in a larger β^n estimate in panel C than in panel A (or B). Moreover, serial correlation is more of a concern in more aggregate data, which again could result in a larger β^n estimate at the broadest level of aggregation in panel C (despite appropriate GLS corrections at each level). Another example is that implicit district weights differ across aggregation levels. For instance, at the most disaggregate level in panel A, different districts in South Wales receive different weights, and each district has a small weight; in contrast, at the most aggregate level in panel C, the whole of South Wales is treated as one single labour market. This could result in a larger β^n estimate in panel C, weighed towards

[13] We use work address for WRS migrants and ASHE workers (to eliminate concerns that they may live in one district and work in another) and home address for JSA claimants, who we assume, search for jobs primarily in their neighbourhood.

South Wales.

4.3 Robustness Checks

The implicit assumption so far is that all WRS migrants compete with all natives in each district, which may not be realistic. This is because the vast majority of WRS migrants do not compete with highly skilled natives. We relax this by assuming that WRS migrants are only substitutes for low-skilled natives (not for high-skilled) within each district. We also experiment with other vulnerable groups, such as female and young natives. Here, the assumption is that WRS migrants are only substitutes for female (young) natives within each district.

By restricting our sample to specific demographic groups we further check the robustness of our earlier estimates. The idea here is that our earlier estimates are for the entire pool of unemployed workers, which might be diluting more adverse effects for low wage workers (Altonji and Card 1991). We thus re-estimate Equation 1 for three groups, in turn: low-skilled (those in elementary occupations), young (those between 18 and 24 years of age) and women. These are workers likely to be competing directly with WRS migrants (see Section 2).

Table 3 shows the associated GLS β^n estimates. Row 1 shows a significant -0.011 estimate for low-skilled workers at the district level (compare with the insignificant 0.012 estimate in row 4 of Panel A of Table 2). This suggests, if anything, a less adverse effect for the low-skilled at the district level. The estimate is less negative, but insignificant, when allowing low-skilled workers to search for jobs at broader aggregation levels. Row 2 shows that for young workers the estimates are again most negative at the broadest aggregation level, but it is never significant. The same is true for female workers. This offers little evidence that migrants are substitutes for low-skilled, young or female natives. Thus, our main conclusion from before of little evidence of adverse claimant unemployment effects is maintained.

4.4 National and Occupational Level

Another way to relax the assumption that all WRS migrants compete with all natives is to assume that low-skilled (high-skilled) WRS migrants compete with low-skilled (high-skilled) natives in a national market. That is, we aggregate the data across occupations (i.e. 9 x 24 cells) and assume that migrants and natives are only substitutes within occupations.

Stratification across occupations – as opposed to stratification across districts – is fruitful because migrants and natives compete more directly within occupations and because bias arising from natives' mobility and migrants' self-selection is less of a concern across occupations. Furthermore, unless natives' mobility and migrants' self-selection bias manifest in exactly the same way across areas and occupations, aggregation across occupations is a further check on the robustness of our earlier estimates.

On the one hand, since WRS migrants are relatively well educated, yet overwhelmingly concentrated into low skilled occupations, this suggests occupational downgrading. This happens when language or labour market barriers prevent migrants to immediately self-select into more favourable occupations (also see Card and DiNardo 2000; Friedberg 2001). Thus, because the WRS inflow was much larger and faster than anticipated, and because it was heavily concentrated into low skilled occupations, concerns about migrants' self-selection bias are reduced. On the other hand, since immediate natives' mobility away from low skilled occupations is limited because it requires retraining (also see Friedberg 2001; Borjas 2003), then these occupations constitute a relatively closed market. (Some limited mobility here derives from occupational progression, which we control for in our regression models.) Thus, because the accession inflow was much larger and faster than anticipated, and because it was heavily concentrated into low skilled occupations, concerns about natives' mobility bias are also reduced.

We therefore re-estimate Equation 1 replacing i with $j = 1,...,9$ to mean occupations (see Section 2.2)[14] and re-defining X_{jt}, due to data limitations, to include the lagged proportion of WRS migrants who are women, young and parents (along with average number of children); their lagged average hours worked; the lagged proportion of unemployed who are women and young; and the lagged average claim duration.

Table 4 shows the associated GLS β^n estimates. Rows 1 to 4 show positive but insignificant estimates: the most complete specification in row 4 shows a 0.035 insignificant estimate. Row 4a shows a negative, though insignificant estimate, when excluding machine operative occupations, where self-selection bias may be a concern. That is because machine operatives may have been hit simultaneously by demand (e.g. booming construction industry) and supply shocks (e.g. WRS migration inflow), as discussed in Section 2.2.

Our results again suggest little evidence of adverse claimant unemployment effects.

[14] Our results here using sought occupation to better capture labour market effects were also robust when we used usual occupation instead.

This is in contrast with Borjas (2006), where more adverse effects were found at wider aggregation levels. Although our results were also successively larger the broader the aggregation level in Table 2, they are smaller at the national level in Table 4 – and they are, if anything, less (not more) adverse at the broadest aggregation level (see Section 4.1).

4.5 Summary

We stratified labour markets in various dimensions (across districts, counties, regions, occupations) and for several demographic groups (low-skilled, young and female) to test alternative assumptions on labour substitutability between migrants and natives. In other words, we considered several alternative local labour markets where migrants might be affecting natives. That is, we modified, in several alternative ways, our assumptions on labour substitutability between migrants and natives. Yet, our estimates were reassuringly small and in the main insignificant across a number of specifications, sub-samples and estimation methods and were not sensitive to the counterfactual underlying each model. [15]

Our main conclusion is that there is little evidence that an increase in the WRS migration rate adversely affected the claimant unemployment rate in Wales between 2004 and 2006. Our results are in line with the international literature, where adverse employment effects are small. They are also in line with the very limited evidence for the UK: Lemos and Portes (2008) reported insignificant claimant unemployment effects when estimating comparable models for the UK using the same sample data. Dustmann et al. (2005) reported insignificant employment and unemployment effects using LFS data for the 1980s and 1990s.

[15] We argue that any remaining endogeneity bias is not very large. Firstly, the WRS migration inflow was a large, rapid, concentrated supply shock resulting mainly from political events. More crucially, the WRS inflow was a shock substantially larger and faster than anticipated, and thus more exogenous than most shocks studied in the literature. As a result, natives' and migrants' responses might have been sufficiently lagged and this reduces concerns of simultaneity bias. Secondly, in the relevant time period, the number of WRS migrants eligible and in receipt of JSA is negligible. Furthermore, our variable of interest is JSA claimant unemployment, as opposed to broader (ILO) unemployment or employment, and this reduces further concerns of simultaneity bias. Thirdly, we used fairly stringent specifications, where we controlled for omitted variables (two of which are natives' mobility and migrants' self-selection) to some extent through district and month-year fixed effects, supply and demand shifters, lagged working age population growth and native netflow rate. Despite controlling for natives' mobility using two alternative proxies, we found little evidence of an associated bias. Fourthly, we found little evidence of an associated bias when we allowed increased natives' mobility through aggregating the data at successively wider levels. Moreover, our results at the national-occupation level, where natives are no longer geographically-bound, showed no evidence of an associated bias. Finally, Lemos and Portes (2008) reported little evidence of bias correction when exploiting a number of carefully defined instruments using the same sample data on comparable models for the UK.

5. Wage Effects

Using a reduced form equation grounded on standard theory (see for example Borjas 1999; Card 2001; Dustmann et al. 2005), we now estimate the effect of the WRS migration inflow on wages in Wales:

$$\Delta W_{iy} = \beta^{w} \Delta M_{iy} + \lambda^{w} \Delta X_{iy} + f_{y}^{w} + \Delta \varepsilon_{iy}^{w} \tag{2}$$

where ΔW_{iy} and ΔM_{iy} are our wage and migration variables, defined in Section 2.2, in district $i = 1,...,22$ and tax-year $y = 1,...,3$; f_{y}^{w} is time fixed effects; ε_{iy}^{w} is the error term; and X_{iy} are labour supply and demand shifters that include the proportion of the total population who are women, young, ethnic minorities and migrants from outside the A8 countries; the lagged proportion of WRS migrants who are women, young and parents (along with average number of children). As before, we estimate Equation 2 in first-difference using GLS and thus district fixed effects were differenced out; time fixed effects are now modelled using year dummies. The interpretation of our coefficient of interest is that a one percentage point increase in the migration rate changes wages by β^{w} %.[16]

Table 5 shows our results across percentiles of the wage distribution. The UK results in Row 4a are again borrowed from Lemos and Portes (2008) and are provided for comparison and completeness, but the main analysis here focuses on the results for Wales. Row 1 of the right-most panel shows a significant 4.214 (unweighted OLS) β^{w} estimate, which corresponds to the raw data in Figure 8. Controlling for district fixed effects produces a 2.745 significant (GLS) estimate, and further controlling for time fixed effects produces a 3.140 significant estimate. Finally, controlling for other supply and demand shocks produces a 3.378 significant estimate. This suggests that an increase of one percentage point in the migration rate raises average wages by 3.4%.

Table 5 shows that the estimates are mostly insignificant below the 50th percentile (see row 4). The estimates for the 60th and 70th percentile are significant in the most complete and preferred specifications in row 4. They suggest that an increase of one percentage point in the migration rate raises wages in the 60th (70th) percentile by 3.9% (5.2%).

As WRS migrants overwhelmingly concentrate around the 5th and 10th percentiles of

[16] One limitation here is that wage data is only available at the yearly level, and as a result, such detailed analysis as the one for claimant unemployment in Sections 3 and 4 was not possible for wages. It is also worth noting that, unlike with the JSA unemployment data, which contained a negligible number of WRS migrants, the ASHE wage data contains both natives and WRS migrants, as discussed in Section 2.1 Thus, it is possible that simultaneity bias, though potentially not too severe, might be more of a concern in our wage models. Our wage estimates were robust, however, when subjected to the same robustness checks to natives' mobility (omitted) variable bias as in Section 4.1.

the wage distribution, we expected to find more adverse (or less favourable) effects there. Our estimates were indeed smaller at the very bottom than higher up the distribution, but they were insignificant. An important point here is that the minimum wage was in force and increasing throughout the period we study, possibly mitigating or offsetting more adverse wage effects for lower paid workers (see Figure 4).[17]

In sum, our main conclusion is that there is little evidence that an increase in the WRS migration rate adversely affected wages in Wales between 2004 and 2006. Our estimates are in line with some evidence in the international literature, where adverse wage effects are small (Grossman 1982; Friedberg 2001; Card 1990 and 2007; Carrasco et al. 2008), though they are in contrast with other evidence of more adverse wage effects (Borjas 1999 and 2006; Angrist and Kugler 2003; Orrenius and Zavodny 2007). They are also in line with the limited evidence available for the UK: Lemos and Portes (2008) reported insignificant wage effects when estimating comparable models for the UK using the same sample data. Using LFS data for the 1980s and 1990s, Dustmann et al. (2005) found no evidence of adverse wage effects and hinted that this may be in part because migrants' skill distribution resembles that of natives. However, Manacorda et al. (2006) argue that the associated relative labour supply change ought to have induced wage effects. Using LFS and BHPS data between the 1970s and 2000s they also found no adverse wage effects and argue that this is because natives and migrants are imperfect substitutes. They then detected some adverse wage effects for earlier migrants. This is in line with findings in Dustmann et al. (2007) of negative wage effects at the bottom of the distribution – where migrants are more concentrated – and positive effects higher up the distribution, when using LFS data for the 1990s and 2000s.

6. Conclusion

The enlargement of the EU in May 2004 triggered a relatively large, rapid and concentrated migration inflow into Wales. We described and evaluated the impact of this inflow on the Welsh labour market. Accession migrants were overwhelmingly concentrated in low-paid low-skilled jobs in elementary occupations and machine operative occupations in the manufacturing and catering sectors. They are concentrated mainly in

[17] Other usual explanations in the literature for insignificant wage and employment effects include factor equalisation as well as industry structure and output mix adjustments. Although neither offers a full explanation (see for example Card 1990 and 2007; Lewis 2003; Ottaviano and Peri 2006), a fruitful avenue for future research is more UK based evidence on both fronts. That would help to understand how the internal flows of goods, capital and labour across markets change following migration inflows and how firms alter their production function and production mix in response to the relative labour supply shift.

cities bordering England – which have been historically more associated with migration – such as Newport, Cardiff, Wrexham and Flintshire.

We found little evidence that the inflow of accession migrants contributed to a fall in wages or a rise in claimant unemployment in Wales between 2004 and 2006. In particular, we found no evidence of an adverse impact on young, female or low-skilled claimant unemployment and no evidence of an adverse impact on the wages of the low-paid. If anything, we found a positive effect on the wages of higher paid workers and some weak evidence of a potentially favourable impact on claimant unemployment. Our results are robust and are in line with other results in the literature. They are also in line with standard theory.

Our unemployment effect estimates were small and in the main insignificant. These estimates were reassuringly robust to a number of specification checks and estimation methods as well as to several different stratifications of the labour market and to different sub-samples of workers.

Our wage effect estimates were positive, small and insignificant at the very bottom of the wage distribution, and were larger higher up, though still insignificant below the median. Estimates for higher paid workers were significant. An increase of one percentage point in the migration rate raises wages of workers in the 60^{th} (70^{th}) percentile of the distribution by 3.9% (5.2%), while it raises average wages by 3.4%.

These results are in line with standard theory, which predicts adverse wages and/or employment effects following a migration inflow that is unbalanced across area or skill. As the accession migration inflow was large, rapid and not balanced across districts or occupations, we expected downward pressure on wages and employment in low-paid low-skilled jobs in occupations and cities where migrants were concentrated. In particular, we expected the wage structure to be affected: competing (complement) workers should have lower (higher) wage increases.

We found evidence that higher paid (complement) workers had larger (positive and significant) wage increases, whereas lower paid (competing) workers had smaller (and insignificantly different from zero) wage increases. One interpretation here is that, relative to higher paid workers, lower paid workers had less favourable (though not adverse) wage increases. Incidentally, more adverse wage effects for lower paid (competing) workers may have been potentially mitigated or offset because they were protected by a concurrently increasing minimum wage.

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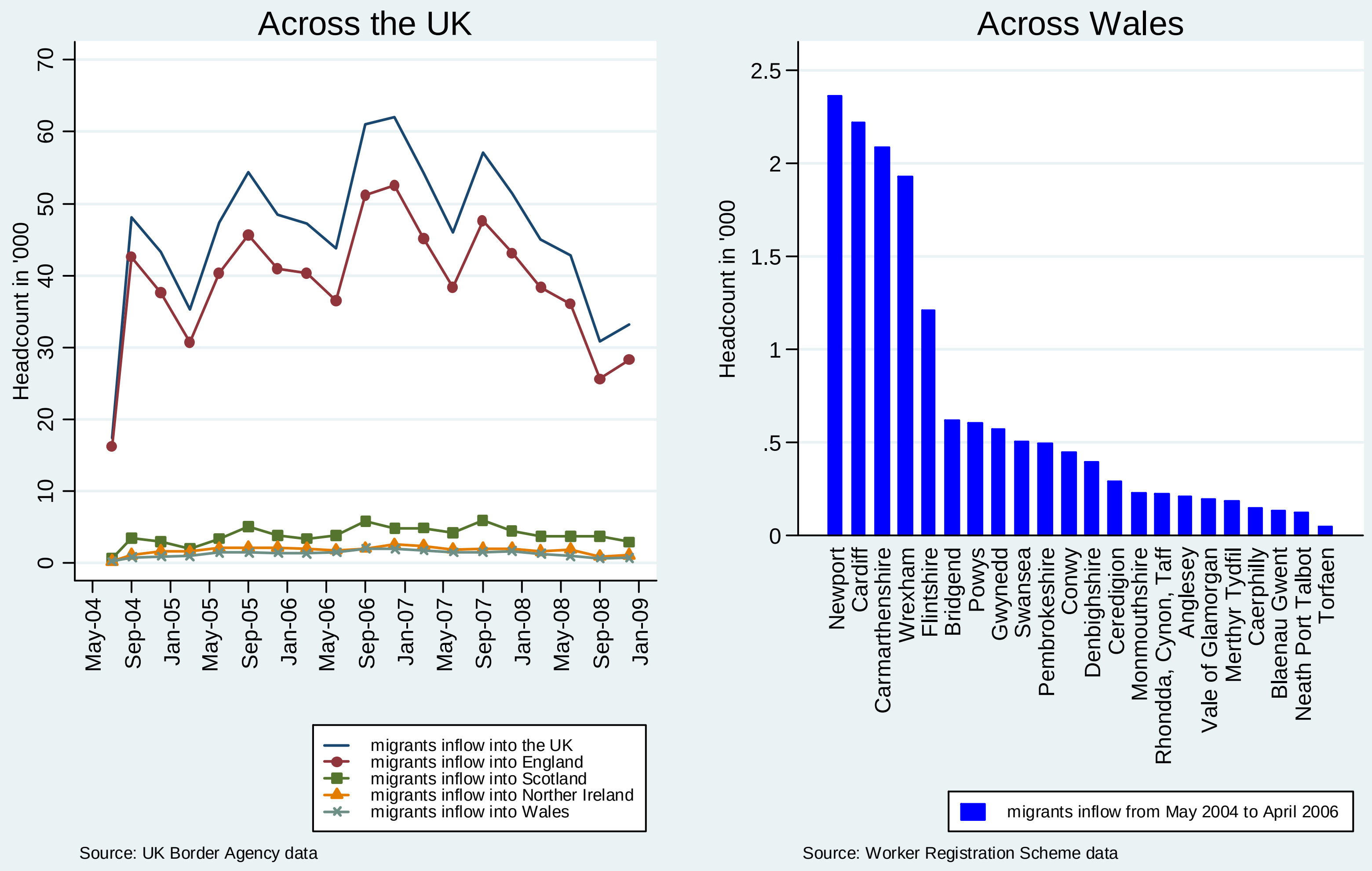
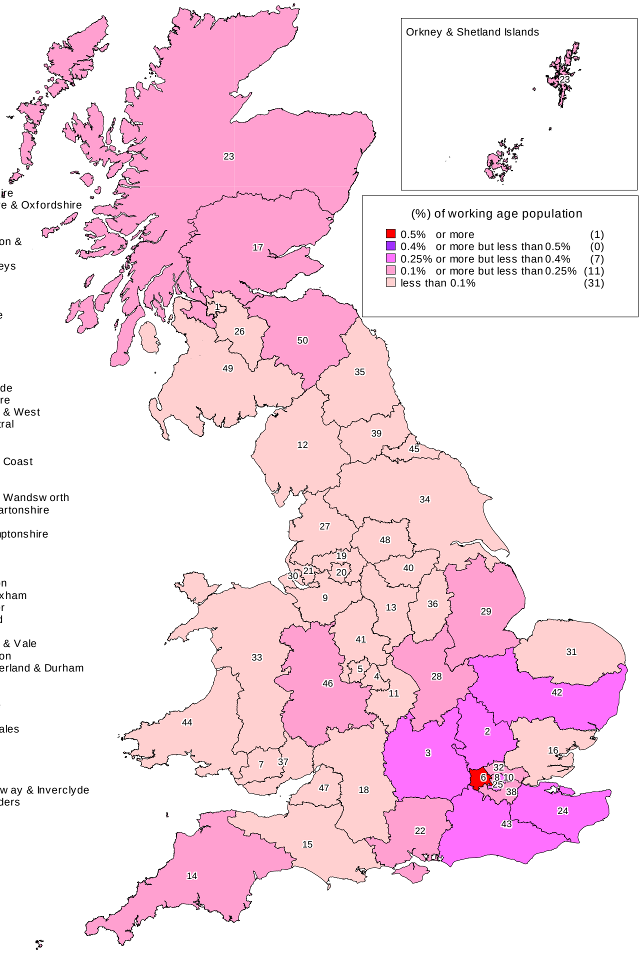
Figure 1 - Migration Inflow



Source: Worker Registration Scheme data and UK Border Agency data

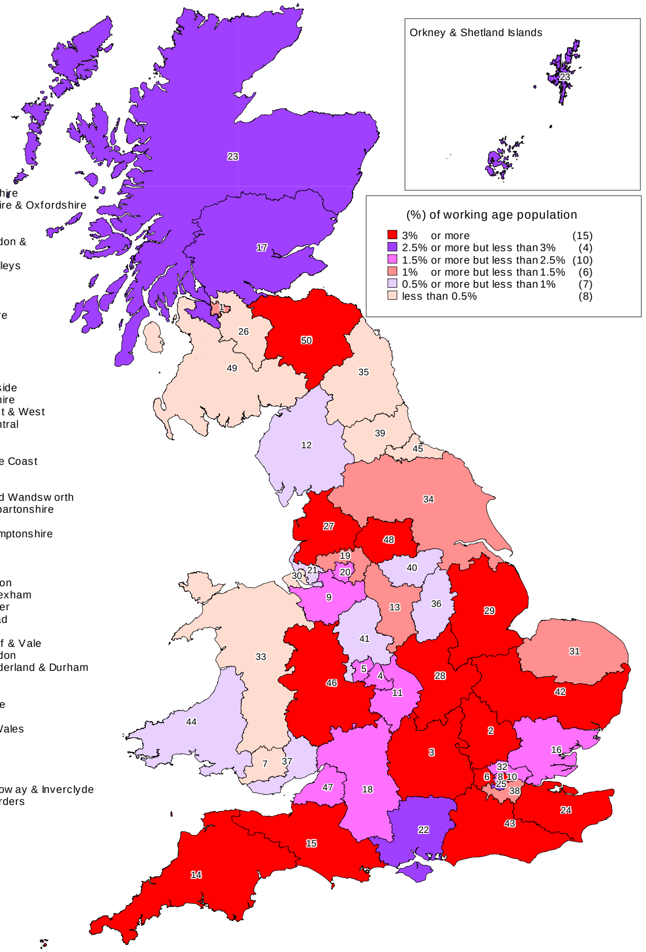
Figure 2 – Migration Inflow Rate by Regions

2004



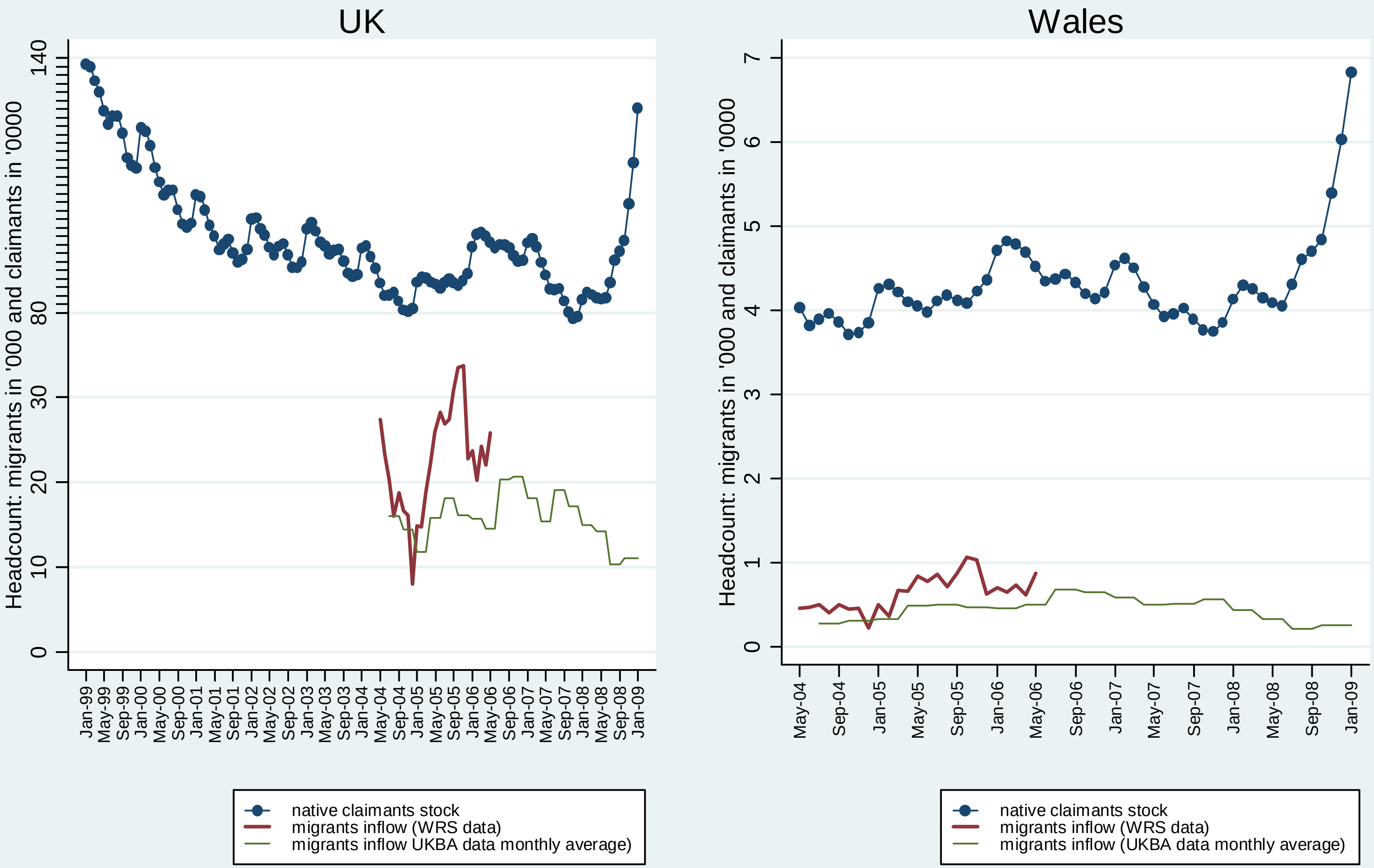
1 Glasgow
2 Bedfordshire & Hertfordshire
3 Berkshire, Buckinghamshire & Oxfordshire
4 Birmingham and Solihull
5 Black Country
6 Brent, Harrow and Hillingdon & West London
7 Bridgend and Eastern Valleys
8 Central London
9 Cheshire
10 City and East London
11 Coventry & Warwickshire
12 Cumbria
13 Derbyshire
14 Devon & Cornwall
15 Dorset & Somerset
16 Essex
17 Forth Valley, Fife & Tayside
18 Gloucestershire & Wiltshire
19 Greater Manchester East & West
20 Greater Manchester Central
21 Greater Mersey
22 Hampshire
23 Highland, Islands & Clyde Coast & Grampian
24 Kent
25 Lambeth, Southwark and Wandsworth
26 Lanarkshire & East Dunbartonshire
27 Lancashire
28 Leicestershire & Northamptonshire
29 Lincolnshire
30 Liverpool & Wirral
31 Norfolk
32 North & North East London
33 North West Wales & Wrexham
34 North Yorkshire & Humber
35 Northumbria & Gateshead
36 Nottinghamshire
37 South East Wales Cardiff & Vale
38 South & South East London
39 South Tyne ,City of Sunderland & Durham
40 South Yorkshire
41 Staffordshire
42 Suffolk & Cambridgeshire
43 Surrey & Sussex
44 Swansea Bay & West Wales
45 Tees Valley
46 The Marches
47 West of England
48 West Yorkshire
49 Ayrshire, Dumfries, Galloway & Inverclyde
50 Edinburgh, Lothian & Borders

2006



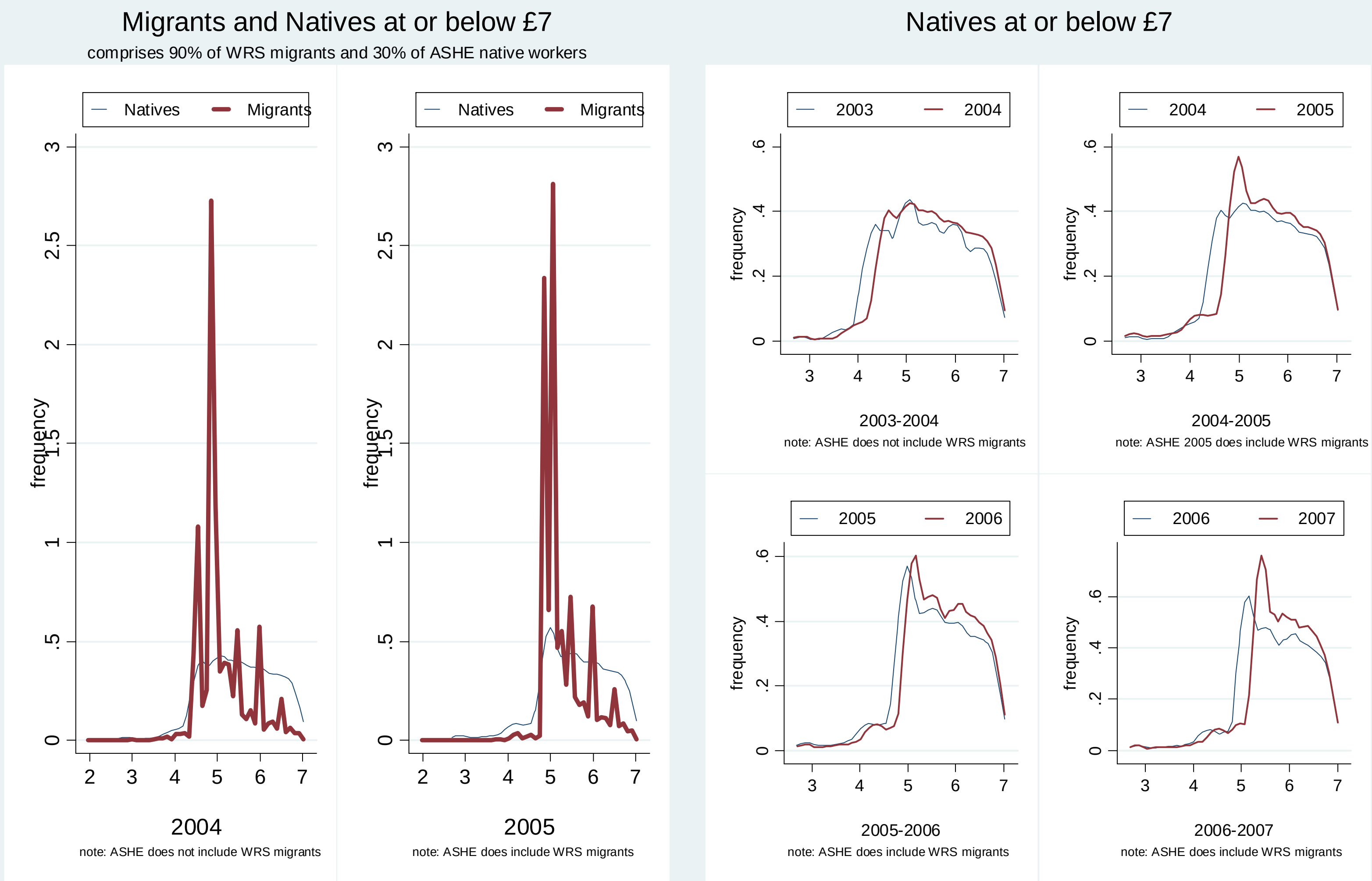
1 Glasgow
2 Bedfordshire & Hertfordshire
3 Berkshire, Buckinghamshire & Oxfordshire
4 Birmingham and Solihull
5 Black Country
6 Brent, Harrow and Hillingdon & West London
7 Bridgend and Eastern Valleys
8 Central London
9 Cheshire
10 City and East London
11 Coventry & Warwickshire
12 Cumbria
13 Derbyshire
14 Devon & Cornwall
15 Dorset & Somerset
16 Essex
17 Forth Valley, Fife & Tayside
18 Gloucestershire & Wiltshire
19 Greater Manchester East & West
20 Greater Manchester Central
21 Greater Mersey
22 Hampshire
23 Highland, Islands & Clyde Coast & Grampian
24 Kent
25 Lambeth, Southwark and Wandsworth
26 Lanarkshire & East Dunbartonshire
27 Lancashire
28 Leicestershire & Northamptonshire
29 Lincolnshire
30 Liverpool & Wirral
31 Norfolk
32 North & North East London
33 North West Wales & Wrexham
34 North Yorkshire & Humber
35 Northumbria & Gateshead
36 Nottinghamshire
37 South East Wales Cardiff & Vale
38 South & South East London
39 South Tyne ,City of Sunderland & Durham
40 South Yorkshire
41 Staffordshire
42 Suffolk & Cambridgeshire
43 Surrey & Sussex
44 Swansea Bay & West Wales
45 Tees Valley
46 The Marches
47 West of England
48 West Yorkshire
49 Ayrshire, Dumfries, Galloway & Inverclyde
50 Edinburgh, Lothian & Borders

Figure 3 - Migration Inflow and Claimant Unemployment



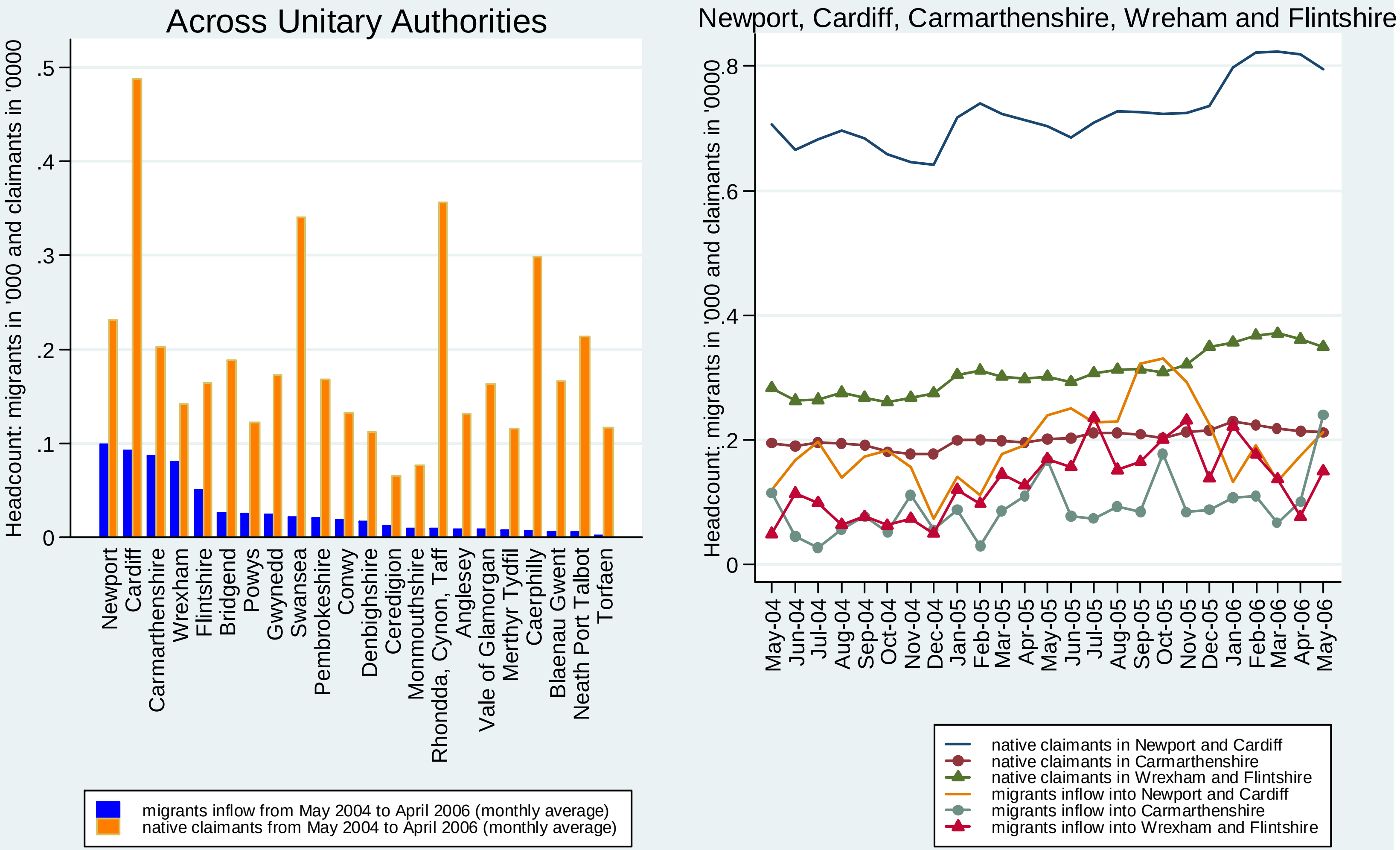
Source: Worker Registration Scheme data, UK Border Agency data and Jobbseeker's Allowance data

Figure 4 - UK Hourly Wage Distributions



Source: Worker Registration Scheme data and Annual Survey of Hours and Earnings data

Figure 5 - Migration Inflow and Claimant Unemployment



Source: Worker Registration Scheme data and Jobbseeker's Allowance data

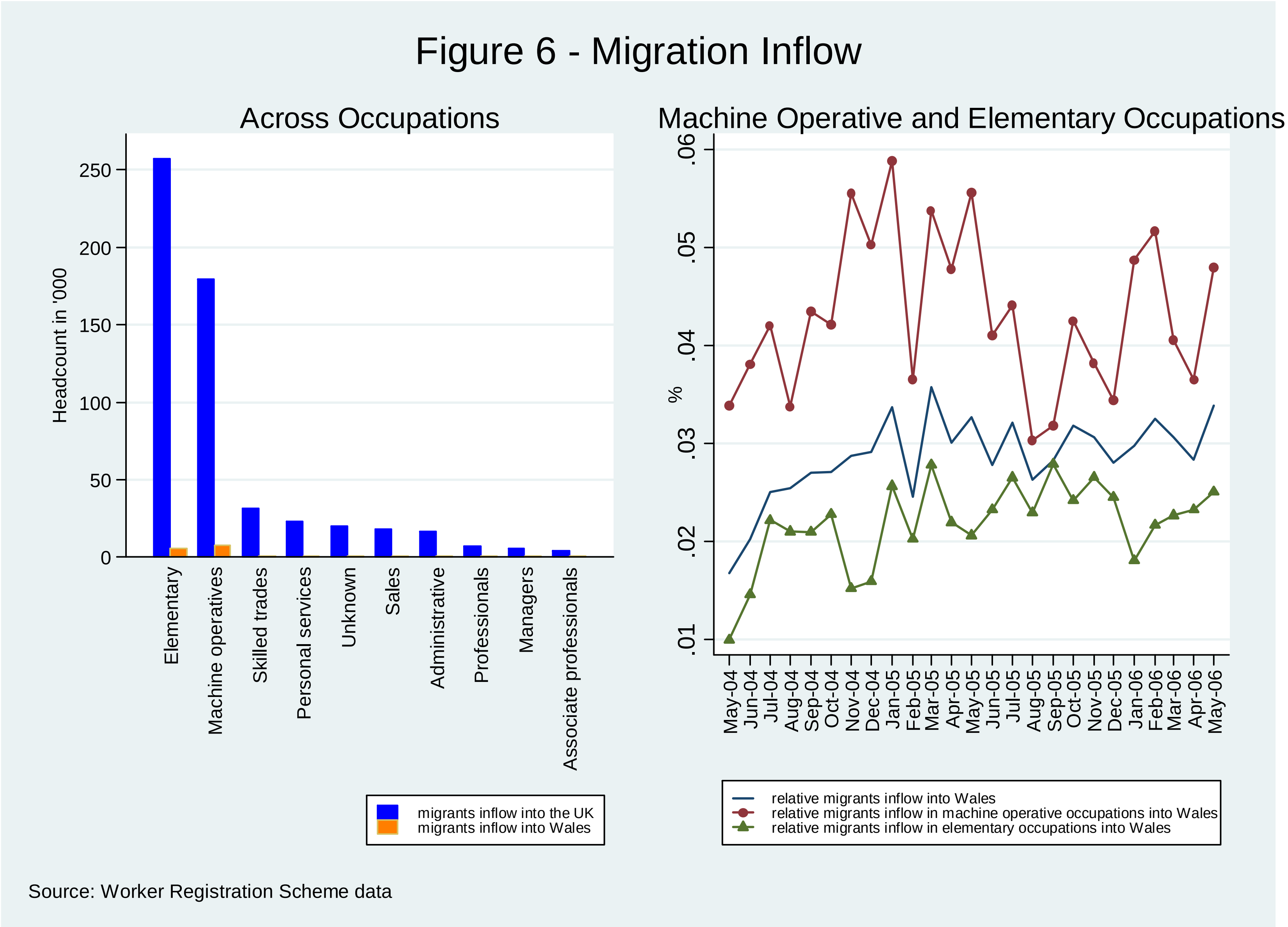
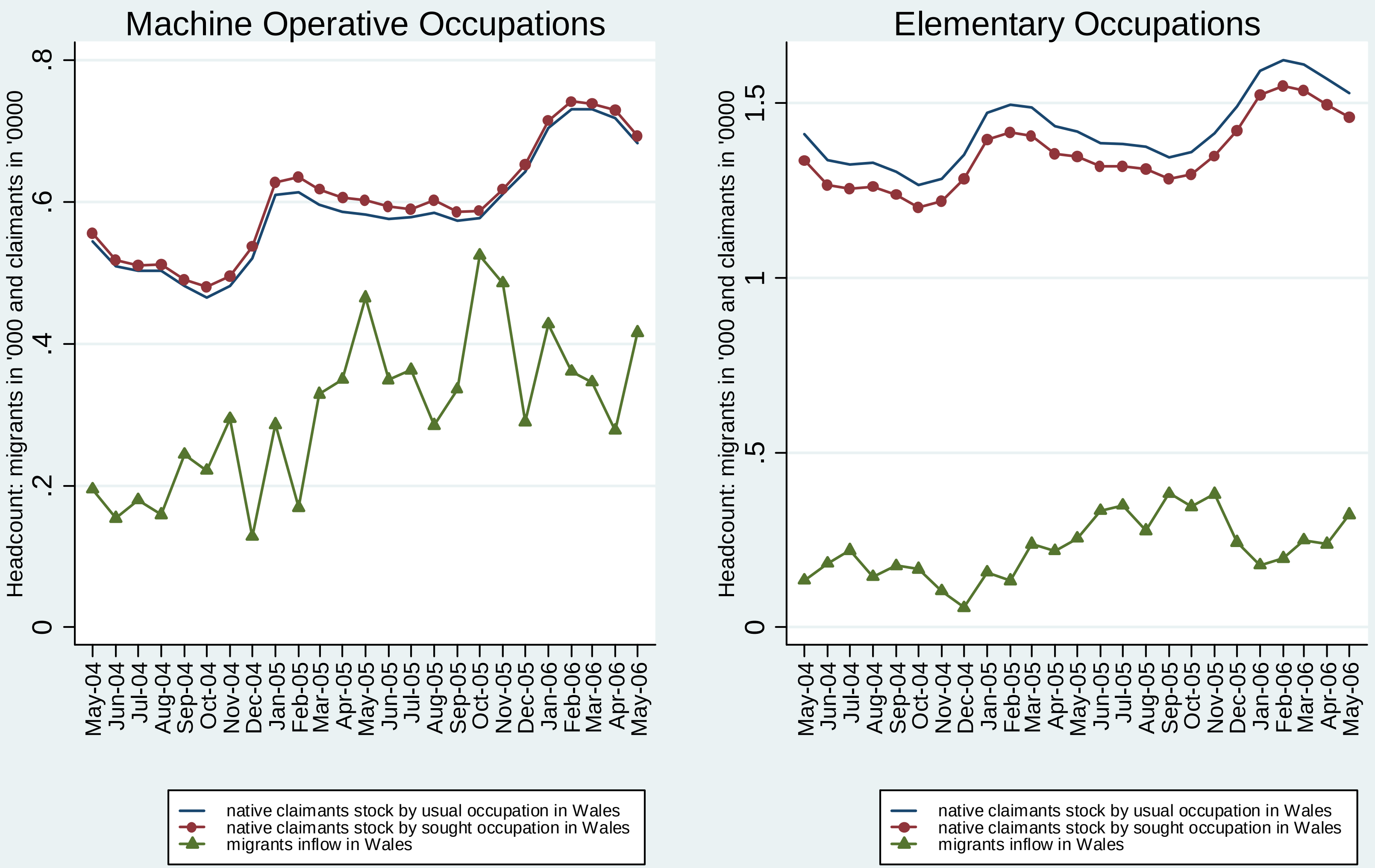


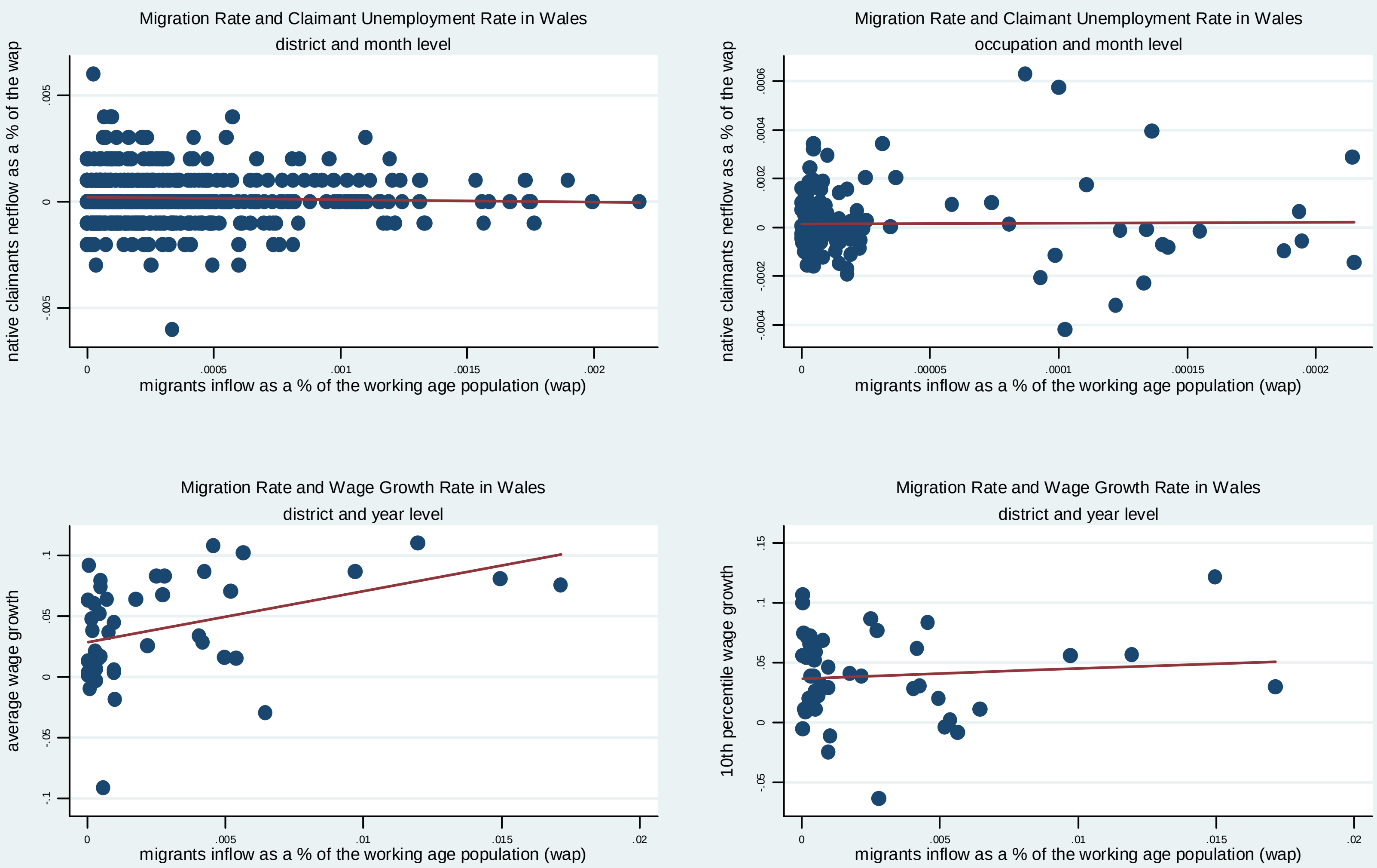
Figure 6 - Migration Inflow
Across Occupations
Headcount in '000
250
200
150
100
50
0
Elementary
Machine operatives
Skilled trades
Personal services
Unknown
Sales
Administrative
Professionals
Managers
Associate professionals
migrants inflow into the UK
migrants inflow into Wales
Machine Operative and Elementary Occupations
%
.06
.05
.04
.03
.02
.01
May-04
Jun-04
Jul-04
Aug-04
Sep-04
Oct-04
Nov-04
Dec-04
Jan-05
Feb-05
Mar-05
Apr-05
May-05
Jun-05
Jul-05
Aug-05
Sep-05
Oct-05
Nov-05
Dec-05
Jan-06
Feb-06
Mar-06
Apr-06
May-06
relative migrants inflow into Wales
relative migrants inflow in machine operative occupations into Wales
relative migrants inflow in elementary occupations into Wales
Source: Worker Registration Scheme data

Figure 7 - Migration Inflow and Claimant Unemployment



Source: Worker Registration Scheme data and Jobbseeker's Allowance data

Figure 8 - Migration Rate, Claimant Unemployment Rate and Wage Growth



Source: Worker Registration Scheme data, Jobbseeker's Allowance data and Annual Survey of Hours and Earnings data

Table 1 - DESCRIPTIVE STATISTICS

VARIABLES	WRS		JSA		ASHE		LFS	
	May 2004 - May 2006		May 2004 - May 2006		May 2004 - May 2006		April 2004 - June 2006	
	migrants		claimants		workers		UK born	Overseas born
I - POPULATION VARIABLES - % of those who are:	**UK**	**Wales**	**UK**	**Wales**				
Aged:								
under 16 years old	0.00	0.00	-	-	na	na	0.21	0.08
16 to 24 years old	0.37	0.37	0.30	0.35	na	na	0.12	0.11
25 to 34 years old	0.45	0.43	0.24	0.23	na	na	0.12	0.24
35 to 64 years old	0.18	0.21	0.45	0.41	na	na	0.40	0.44
over 65 years old	0.00	0.00	0.00	0.00	na	na	0.16	0.13
Women	0.43	0.38	0.74	0.76	na	na	0.51	0.52
Parents (with dependent children)	0.06	0.06	na	na	na	na	0.27	0.32
Blacks	-	-	na	na	na	na	0.01	0.12
Asians	-	-	na	na	na	na	0.02	0.25
Nationality:								
Polish	0.61	0.67	na	na	na	na	-	0.02
Lithuanian	0.12	0.08	na	na	na	na	-	0.01
Slovakian	0.10	0.13	na	na	na	na	-	0.00
Lativian	0.07	0.03	na	na	na	na	-	0.00
Located in:								
London	0.17	-	0.19	-	na	na	0.09	0.41
South East	0.14	-	0.08	-	na	na	0.14	0.13
East of England	0.12	-	0.07	-	na	na	0.09	0.08
East Midlands	0.09	-	0.06	-	na	na	0.07	0.05
Yorkshire and the Humber	0.08	-	0.09	-	na	na	0.09	0.06
West Midlands	0.08	-	0.11	-	na	na	0.09	0.07
North West	0.08	-	0.12	-	na	na	0.12	0.07
South West	0.08	-	0.05	-	na	na	0.09	0.05
Scotland	0.08	-	0.10	-	na	na	0.09	0.04
Northern Ireland	0.04	-	0.03	-	na	na	0.03	0.01
Wales	0.03	-	0.05	-	na	na	0.05	0.02
North East	0.01	-	0.05	-	na	na	0.05	0.02
II - LABOUR MARKET VARIABLES - % of those who are in:								
Occupations:								
elementary occupations	0.46	0.36	0.35	0.34	na	na	0.11	0.14
machine operatives occupations	0.32	0.49	0.10	0.14	na	na	0.08	0.07
skilled trades occupations	0.06	0.05	0.11	0.12	na	na	0.12	0.08
personal services occupations	0.04	0.04	0.05	0.05	na	na	0.08	0.08
unknown occupation	0.04	0.00	0.01	0.01	na	na	0.00	0.00
sales and customer service occupations	0.03	0.02	0.13	0.12	na	na	0.08	0.07
administrative occupations	0.03	0.02	0.10	0.10	na	na	0.13	0.09
professional occupations	0.01	0.01	0.04	0.00	na	na	0.12	0.17
managers and senior officials	0.01	0.01	0.04	0.03	na	na	0.15	0.15
technical occupations	0.01	0.01	0.06	0.06	na	na	0.14	0.15
Sectors:								
manufacturing	0.31	0.48	na	na	na	na	0.13	0.11
distribution, hotels & restaurants	0.27	0.23	na	na	na	na	0.19	0.21
transport & communication	0.09	0.07	na	na	na	na	0.07	0.08
agriculture and Fishing	0.08	0.02	na	na	na	na	0.01	0.01
banking, finance & insurance etc	0.08	0.06	na	na	na	na	0.15	0.19
public admin, educ & health	0.06	0.06	na	na	na	na	0.28	0.28
construction	0.04	0.03	na	na	na	na	0.08	0.05
other services	0.02	0.02	na	na	na	na	0.06	0.06
energy and water	0.00	0.00	na	na	na	na	0.01	0.01
Part time	0.08	0.05	na	na	na	na	0.26	0.22
Employment rate	-	-	-	-	na	na	0.76	0.67
Unemployment rate	-	-	-	-	na	na	0.05	0.07
Average claim duration	-	-	31.32	29.83	na	na	-	-
Looking for a job in their usual occupation	-	-	0.84	0.87	na	na	-	-
Average hours worked	37.83	38.77	-	-	na	na	36.87	38.37
					2004	**2006**	**April 2004 - March 2006**	
5th percentile hourly wage distribution	4.50	4.50	-	-	4.77	5.16	4.50	4.61
10th percentile hourly wage distribution	4.65	4.69	-	-	5.14	5.55	5.30	5.18
20th percentile hourly wage distribution	4.85	4.85	-	-	5.99	6.45	6.32	6.25
30th percentile hourly wage distribution	4.87	4.85	-	-	6.92	7.45	7.27	7.45
40th percentile hourly wage distribution	5.00	5.00	-	-	7.95	8.55	8.26	8.61
50th percentile hourly wage distribution	5.05	5.05	-	-	9.18	9.89	9.40	9.89
60th percentile hourly wage distribution	5.20	5.05	-	-	10.75	11.63	10.87	11.54
70th percentile hourly wage distribution	5.50	5.20	-	-	12.80	13.89	12.66	13.57
80th percentile hourly wage distribution	6.00	5.50	-	-	15.56	16.92	15.18	16.56
90th percentile hourly wage distribution	6.73	6.20	-	-	20.47	22.29	19.24	21.63
Average hourly wage distribution	5.56	5.32	-	-	12.04	13.09	11.31	12.18
Standard deviation hourly wage distribution	2.03	1.66	-	-	na	na	6.96	8.10
Adult minimum wage	4.80	4.80	-	-	4.50	5.05	4.80	4.80
number of observations	562830	16137	22016120	1039123	21915	23725	201294305	21169990

Source: Worker Registration Scheme data, Jobseeker's Allowance data, Annual Survey of Hours and Earnings and Labour Force Survey

(1) Variables not available or not defined in a particular dataset are indicated by "na" or "-". For example, the employment and unemployment rates are not definied for the WRS ASHE or JSA, where all individuals are working/unemployed. The proportion of parents from the LFS is for 2006 Q2, where the household weight used is based on 2003 population estimates as re-weighted household datasets are yet unavailable (the other figures are based on 2007 population estimates).

(2) As ASHE is not available at the micro level, we are unable to compute percentiles for the period 2004-2006; we instead report percentiles for 2004 and 2006 directly from the "ASHE tables" available from the ONS. Similarly, standard deviation is not available.

(3) As detailed in the text (see Section 2), the WRS measures inflows, whereas the JSA and LFS measure stocks. Therefore, the WRS figures are cumulative.

(4) National minimum wage (adult rate) is: £4.50 between 1 October 2003 and 30 September 2004; £4.85 between 1 October 2004 and 30 September 2005; £5.05 between 1 October 2005 and 30 September 2006.

Table 2 - UNEMPLOYMENT EFFECTS OF MIGRATION

Models	coefficient	s. errors
A - Twenty-Two-Way District Aggregation		
(1) Raw coefficient	**-0.115**	0.140
(2) Baseline specification	**0.024**	0.035
(3) Adding time effects	**0.006**	0.028
(4) Adding demand and supply controls	**0.012**	0.033
(5) Adding working age population growth	**0.024**	0.032
(6) Adding native netflow rate	**0.017**	0.032
(6a) UK	***0.003***	*0.078*
B - Seven-Way District Aggregation		
(1) Raw coefficient	**-0.052**	0.261
(2) Baseline specification	**-0.038**	0.173
(3) Adding time effects	**0.010**	0.104
(4) Adding demand and supply controls	**-0.068**	0.168
(5) Adding working age population growth	**-0.137**	0.214
(6) Adding native netflow rate	**-0.143**	0.219
(6a) UK	***0.057***	*0.086*
C - Three-Way District Aggregation		
(1) Raw coefficient	**-0.227**	0.558
(2) Baseline specification	**-0.465**	0.490
(3) Adding time effects	**-0.539**	0.465
(4) Adding demand and supply controls	**-1.522**	0.739
(5) Adding working age population growth	**-1.522**	0.731
(6) Adding native netflow rate	**-1.439**	0.721
(6a) UK	***0.115***	*0.106*

(a) These are GLS estimates weighted by the sample size used to calculate the dependent variable (except in row 1, where OLS unweighted estimates are shown).
(b) The dependent variable is the claimant unemployment rate and the independent variable of interest is the WRS migration rate (see Sections 3 and 4).
(c) Time fixed effects are modeled with month dummies; area fixed effects are differenced out. See Section 3 for discussion on demand and supply controls.
(d) The interpretation of the coefficient is that a 1 percentage point increase in the WRS migration rate changes the claimant unemployment rate by b percentage points.
(e) The estimates for the UK in row 6a of each panel are borrowed from Lemos and Portes (2008). The number of districts for the UK in panel A row 6a is 409 (i.e. 409 x 24 observations), the number of counties for the UK in panel B row 6a is 49 (i.e. 49 x 24 observations) and the number of government regions (one of which is Wales) for the UK in panel C row 6a is 12 (i.e. 12 x 24 observations). The number of observations for Wales in each panel is respectively 22 x 24, 7 x 24 and 3 x 24. Thus, the estimates for Wales in rows 1-6 and the estimates for the UK in row 6a of each panel are not directly comparable, though they follow the same pattern of ever broader aggregation.

Table 3 - UNEMPLOYMENT EFFECTS OF MIGRATION (robustness checks)

Models	coefficient	s errors
A - Twenty-Two-Way District Aggregation		
(1) Low Skilled	**-0.011**	0.005
(2) Young	**-0.024**	0.016
(3) Female	**-0.010**	0.008
B - Seven-Way District Aggregation		
(1) Low Skilled	**-0.003**	0.056
(2) Young	**0.020**	0.088
(3) Female	**-0.023**	0.061
C - Three-Way District Aggregation		
(1) Low Skilled	**-0.413**	0.256
(2) Young	**-0.476**	0.328
(3) Female	**-0.398**	0.298

(a) Notes as in Table 2.
(b) All estimates here to be compared with estimates in row (4) of each respective panel of Table 2.

Table 4 - UNEMPLOYMENT EFFECTS OF MIGRATION (by occupation)

Models	coefficient	s. errors
(1) Raw coefficient	**0.141**	0.139
(2) Baseline specification	**0.097**	0.114
(3) Adding time effects	**0.024**	0.120
(4) Adding demand and supply controls	**0.035**	0.064
(4a) Excluding machine operative occupations	**-0.327**	0.189
(4b) UK Excluding machine operative occupations	***-0.049***	*0.089*

(a) Notes as in Table 2, except that the number of observations is now 9 x 24. As before, time fixed effects are modeled with month dummies. Occupation fixed effects are differenced out.

Table 5 - WAGE EFFECTS OF MIGRATION

Models	5th percentile		10th percentile		20th percentile		30th percentile		40th percentile	
	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors
Twenty-Two-Way District Aggregation										
(1) Raw coefficient	**1.271**	1.796	**0.851**	1.419	**0.533**	1.446	**0.850**	1.529	**2.608**	1.713
(2) Baseline specification	**1.053**	1.226	**0.247**	1.550	**-0.259**	1.852	**0.261**	1.864	**1.540**	1.989
(3) Adding time effects	**1.032**	1.467	**2.619**	2.018	**2.382**	2.044	**3.213**	2.011	**3.983**	2.060
(4) Adding demand and supply controls	**1.360**	1.429	**2.845**	2.190	**2.410**	2.294	**3.079**	2.181	**4.175**	2.343
(4a) UK	**0.212**	0.190	**0.110**	0.220	**0.162**	0.305	**0.365**	0.239	**0.453**	0.250

Models	50th percentile		60th percentile		70th percentile		80th percentile		90th percentile		Average wage	
	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors	coefficient	s. errors
(1) Raw coefficient	**3.254**	1.702	**2.437**	1.690	**4.715**	2.059	**-**	-	**-**	-	**4.214**	1.481
(2) Baseline specification	**2.128**	2.216	**1.099**	1.881	**3.777**	1.763	**-**	-	**-**	-	**2.745**	1.463
(3) Adding time effects	**4.973**	2.602	**3.957**	1.312	**5.486**	1.908	**-**	-	**-**	-	**3.140**	1.210
(4) Adding demand and supply controls	**5.004**	2.735	**3.874**	1.497	**5.222**	2.316	**-**	-	**-**	-	**3.378**	1.254
(4a) UK	**0.438**	0.307	**0.455**	0.309	**0.460**	0.336	**0.586**	0.410	**0.869**	1.743	**0.246**	0.276

(a) Notes as in Table 2, except that the dependent variable is now the average and various percentiles of the wage distribution across years and districts, and that the number of observations is now 22 x 3.
(b) Estimates not available are indicated by "-". This is due to small sample size and/or non-reliability or non-availability of data points, as explained in detail in the "ASHE tables" available from the ONS. Even though where estimates for Wales are missing corresponding estimates for the UK are reported, care should be taken as these suffer the same limitation: they are based on substantially smaller sample size (missing data points) and are here reported for completeness only.
(c) The estimates for the UK in row 4a of each panel are borrowed from Lemos and Portes (2008), where the number of districts is 409 (i.e. 409 x 3 observations).