

Abstract

This paper considers three aspects of the job insecurity facing British men in the last two decades. The probability of becoming unemployed, the costs of unemployment in terms of real wages losses and the probability that the continuously employed will experience substantial real wage losses. The first of these has not risen in the last two decades, the second has risen by around 50 percent and the third has risen, particularly for the top skill groups.

This paper was produced as part of the Centre's
Labour Markets Programme

A Picture of Job Insecurity Facing British Men

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November 2000

Published by
Centre for Economic Performance
London School of Economics and Political Science
Houghton Street
London WC2A 2AE

© Stephen Nickell, Tracy Jones and Glenda Quintini, submitted August 1999

ISBN 0 7530 1423 8

Individual copy price: £5

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Introduction	1
2. The Chances of Becoming Unemployed	2
3. The Wage Losses Consequent on Unemployment	4
4. How Common are Falls in Real Wages?	8
5. Conclusions	11
Tables	13
Figures	26
References	28

Acknowledgements

We are most grateful to the Leverhulme Trust for funding under the auspices of the programme on the Labour Market Consequences of Technological and Structural Change. Glenda Quintini is funded by the TMR programme of the European Union. We are grateful to Richard Freeman, Larry Katz, Richard Dickens, Peter Sloane, Chris Pissarides and participants at seminars in Harvard, Queens University, Belfast, Bonn and Aberdeen for helpful comments on an earlier version.

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Introduction

It is now a commonplace view in Britain that job insecurity has risen significantly over the last two decades. Yet providing evidence of substantial changes in the job market that supports this view has not proved easy. Data on job tenure, for example, exhibits no dramatic changes. Burgess and Rees (1996) examine various aspects of job tenure using the UK General Household Survey and find that average elapsed tenure for men fell from around 10.5 years in the mid 1970s to around 9.4 years in the early 1990s. There has been no noticeable change for women over the same period. This picture is consistent with that reported by Gregg *et al* (1997) using the UK Labour Force Survey, who find that median elapsed job tenure for men has fallen slowly but steadily from 1975 to 1995 with the overall fall being 20 percent over the whole period. Again, for women, the change is much less significant, in part because of the increasing number of women who do not take a formal job break when having children.

How do these apparently rather small shifts compare with the opinions of the workers themselves? In the OECD's systematic analysis of this question in Chapter 5 of the 1997 Employment Outlook, they report a significant rise in the proportion of both men and women who are not completely satisfied with job security from 61.7 percent in 1991 to 78.4 percent in 1995, with all the increase coming between 1991 and 1992. These numbers from the British Household Panel Study appear, in fact, to be a statistical artefact arising from a change in the "showcard" used when asking the relevant question (see Green *et al* 1998, p.6, for precise details). However, the OECD also reports a massive fall of 22 percentage points in the proportion of employees who respond favourably on the job security aspects of their work between 1985 and 1995. Green *et al* (1998) are also sceptical about this result, pointing out that the data on which it is based, collected by International Survey Research (ISR) Ltd, are not generated by anything close to a random sample but are based on the workforces of ISR clients in any particular year. Green *et al* (1998) then report, on the basis of data collected by the ESRC Social Change and Economic Life Initiative in 1986 and the Skills Survey in 1997, that the average reported expected risk of job loss has changed little over the relevant period, although it has risen for professional workers.

At present, therefore, the overall picture of changes in job security is somewhat cloudy. This is, in part, due to the fact that the objective data on job tenure are not very informative about feelings of insecurity. Individuals feel insecure at work when there is a significant probability that they will become substantially worse off. This may occur in a variety of ways. They may feel insecure because there is a high probability that they will lose their job. However, this feeling of insecurity will surely be exacerbated if the cost of losing their job is also high. Thus we know, for example, that the quality of re-entry jobs after unemployment has fallen substantially from the 1970s to the 1990s (see Gregg and Wadsworth, 1996), so it is possible that the cost of job loss has risen over the same period¹.

¹ The decline in quality of re-entry jobs does not necessarily indicate a rise in the cost of job loss because exit jobs may also have declined in quality in precisely the same fashion.

But feelings of insecurity may not only be related to job loss. Such feelings can also be engendered by a high probability that real wages will fall substantially in a continuing job. Indeed insecurity can rise in a world where jobs remain secure precisely because wages have become more “flexible”. So our aim in what follows is to present a picture of recent changes in the chances of becoming unemployed, (Section 1), in the wage consequences of becoming unemployed (Section 2) and in the chances of significant real wage falls in continuing jobs (Section 3). We also focus particularly on men, essentially because the existing evidence suggests that men are likely to have seen more substantial changes in job insecurity than women over the last two decades. Our findings are summarised in the Conclusions.

2. The Chances of Becoming Unemployed

One of the main fears of the employed is that of losing their job. But not everyone who leaves a job and becomes unemployed does so involuntarily. Some people resign their jobs to enter unemployment and, presumably, they feel they are better off by doing so². Since we wish to interpret a rise in the chances of an employee becoming unemployed as corresponding to a rise in insecurity, we must first check to see if the proportion who leave their jobs and become unemployed voluntarily has been subject to any systematic shifts, for such shifts might corrupt our desired interpretation of the numbers. Luckily, as Figure 1 makes clear, the proportion of the unemployed who resigned from their previous jobs exhibits no trend³, at least since 1981. In particular, we find that for the three periods considered subsequently, namely 1982-86, 1987-91, 1992-97, the proportion of unemployed who resigned is 10.1, 11.2 and 10.0 percent respectively. So looking at all the unemployed is not going to generate misleading results when we are concentrating on changes over time.

Unemployment entry probabilities

Since the 1960s, unemployment rates among men in Britain have risen dramatically and even today, when unemployment is lower than it has been for many years, it is still more than twice what it was in the late 1960s (see Figure 2a). Despite this, the probability of an employed man in Britain entering unemployment is actually lower today than in the late 1960s (see Figure 2b), although earlier in the 1990s this probability had attained unprecedented heights. So while there is no obvious secular trend in the chances of a working man becoming unemployed, there is some evidence that when the economy is entering a serious slump as in the mid 1970s, the early 1980s and the early 1990s, the chances of becoming unemployed have tended to be higher in the most recent episode. Unfortunately, it is not possible to break down the unemployment inflow data by any variables of interest, notably by occupation or skill group. In light of this, we next pursue the issue of entry probabilities by looking at the probability of a worker being unemployed twelve months later.

² Of course, even individuals who resign may have effectively been sacked. Employees who antagonise their bosses are often asked to resign or placed in a position where they have no alternative but to resign.

³ There are other categories where individuals leave their jobs and enter unemployment without any necessary interference from the employee; for example, leaving because of sickness or for family reasons. In fact, these proportions have also remained stable. In any event, it is arguable that entering unemployment for these reasons is part of job insecurity.

The chances of a male worker being unemployed twelve months hence

An alternative measure of insecurity related to job loss is to ask the question, what are the chances of an employed man being unemployed twelve months later? This depends on both the probability of entry *and* the duration of the unemployment spell, so it will depend to some extent on the unemployment rate. It is obviously true that even if there is no secular increase in the probability of entering unemployment, a systematically higher unemployment rate will raise insecurity because the unemployment spell entered into will have a longer expected duration. The expected length of the unemployment spell is a part of the cost of job loss, so we are simply arguing here that insecurity is increasing both in the probability of job loss and its expected cost⁴. Following on from this, we can see from Figure 2a that changes in aggregate unemployment cannot have contributed directly to any secular increase in insecurity since the early 1980s because the trend level of unemployment has not increased since that time, indeed if anything it has declined.

So what we are interested in here is whether there is any evidence that, either overall or for certain groups, the probability of an employee being out of work in twelve months time has risen systematically, once we control for the direct impact of the aggregate unemployment rate.

To pursue this issue we make use of the UK Labour Force Survey. For each year of the survey, t , we compute the probability that a man who was employed in a particular socioeconomic (SEG) or occupation (KOS) group⁵, i , at a certain time in year $t-1$ was unemployed twelve months later. We measure this probability, p_{it} say, by taking the sample proportion. Each year the Survey contains around 70,000 employed men divided into 7 SEGs or 16 KOS groups. So our measured probabilities, which are typically around 4 percent, have a relatively small average sampling standard error of around 0.3 percentage points in the case of KOS groups and around 0.2 percentage points in the case of SEGs⁶, using the standard formula $\left([p(1-p)/\text{sample size}]^{1/2}\right)$.

We then take these sample probabilities and using SEG or KOS groups over time as the unit of observation, we use them to run (fixed effects) panel regressions of the form

$$p_{it} = \mathbf{a}_0 + \mathbf{g}u_t + \sum_{i=2}^n \mathbf{a}_i d_i + \sum_{i=1}^n \mathbf{b}_i d_i f(t) + \mathbf{e}_{it} \quad (1)$$

$i=1 \dots n$ (SEG or KOS), $t=1 \dots T$ (77,79, 81, 83, 84-90 for KOS;
79, 81 83, 84-98 for SEG)

p_{it} = probability that a man who was employed in $t-1$ is unemployed 12 months later

⁴ Of course, it may be argued that if the increased expected spell duration comes about because of a rise in the level of unemployment benefit, this will not be associated with a rise in the cost of job loss. As it happens, using the OECD summary measure of the benefit replacement ration (see OECD Jobs Study, 1994, Table 8.1), we find that this has been falling in Britain since the late 1970s, basically because over most of this period, benefits have been indexed to prices and have therefore grown systematically more slowly than earnings as real wages have risen (see Nickell and van Ours, 1999, Figure 9, for the precise numbers).

⁵ The SEGs are reported in Table 1. The KOS groups are KOS1: Professional supporting management; KOS2: Professional in education, health, welfare; KOS3: Literary, artistic, sport; KOS4: Professional in science and technology; KOS5: Managerial; KOS6: clerical; KOS7: Selling; KOS8: Security, protective services; KOS9: Catering, cleaning, hairdressing, other services; KOS10: Farming, fishing; KOS11: Materials processing (exc. Metal, electrical); KOS12: Materials processing (metal, electrical); KOS13: Painting, repetitive assembling, packaging; KOS14: Construction, mining; KOS15: Transport; KOS16: Miscellaneous.

⁶ Of course, the KOS groups and the SEGs are not of uniform size but even the proportion in the smallest group has a sampling standard error that is only around one-fifth of its size. Overall, given the proportions are used as the dependent variable, there is easily enough "true" variation in these data to be able to detect important trends if they are present in reality.

u_t = aggregate unemployment rate,
 d_i = SEG or KOS dummies,
 $f(t)$ = linear trend or grouped time dummies

The idea here is to see if there have been significant increases over time in the probability of ending up unemployed for men in any particular occupation groups, once we control for the aggregate unemployment rate⁷.

In Table 1 we present an example of a regression based on socioeconomic groups. As we can see, while fluctuations in aggregate unemployment have a strong impact on the probability of an employee being unemployed one year later, there is no evidence whatever of any socioeconomic group exhibiting any systematic rise in this probability over and above this aggregate unemployment effect. Furthermore, if we replace the linear trend by a series of three steps, the same results ensue. Indeed, in all socioeconomic groups, the last step (1991-98) turns out to be at the same level or lower than the first step (1979-84) except for SEG2 where it is about one standard error above. Finally, exactly the same picture applies if we replace the SEGs by the more refined KOS groupings.

Summary

If the insecurity of male employees in Britain has risen over the last eighteen years it has not done so because they are more likely to become unemployed. There has been no systematic increase in the chances of becoming unemployed either on average or within any particular occupational or socioeconomic group. The only significant change is that during the recession of the early 1990s, the probability of unemployment entry rose to a significantly higher level than its maximum during the previous two recessions. On average, however, this probability has exhibited no change. Our next step will be to pursue the issue of the costs of job loss.

3. The Wage Losses Consequent on Unemployment

It is a well known fact that, for one reason or another, workers who lose their jobs and have a spell of unemployment tend to return to work at a lower rate of pay and often suffer a permanent pay reduction (see, for example, Chowdhury and Nickell, 1985; Addison and Portugal, 1989; Swain and Podgursky, 1991; Ruhm, 1991; Jacobson *et al.* 1993 for the US; Gregory and Jukes, 1997 for the UK). Our aim in this section is to investigate the hypothesis that these wage reductions have got bigger since the early 1980s. Furthermore, we would like to interpret any such increase as signifying that individuals who become unemployed have become worse off, thereby corresponding to an increase in insecurity even if the chances of becoming unemployed have not risen.

Our method of investigation is to use fixed effects earnings regressions over three sample periods 1982-86, 1987-91, 1992-97 and see if the *ceteris paribus* (negative) impact of an unemployment spell on earnings has increased in absolute size from the first sample period to the last. Although we used fixed effects regressions, some care is required in the interpretation of our results. First, all we observe is that the individuals concerned pass through an unemployment spell. While we know the length of the spell, we have no information as to why they became unemployed. Indeed, some may have chosen to resign

⁷ If (1) is estimated using the log odds ration, $\ln(p_{it} / (1 - p_{it}))$, as the dependent variable, the pattern of results is identical.

their job and enter unemployment. So the earnings losses we observe are averaged over individuals who enter unemployment for different reasons. Presumably individuals who resign from their previous job to become unemployed are likely to suffer smaller earnings losses than those who lose their previous job involuntarily. This we see as less of a problem than it might be because the proportion of the unemployed who resign from their previous jobs is both small and trendless (see Figure 1).

A second issue concerns whether or not an increased earnings loss genuinely corresponds to an increase in insecurity. To pursue this question, we must consider why the earnings losses arise in the first place and why such losses may have increased. If they arise because of discrimination against the unemployed⁸ and they have increased because of increased discrimination, this obviously corresponds to a rise in insecurity. Suppose, however, that the earnings losses increase because human capital is becoming more specific. Given a constant probability of entry into unemployment, this would represent a rise in insecurity which we would expect to be more marked at higher skill levels, where specific human capital is more likely to be important.

Alternatively, earnings losses can arise as firms weed out incompetent employees who, because of wage uniformity on specific jobs, are earning “too much”. The earnings losses then reflect the extent of this excess pay as their subsequent wages are, on average, closer to the correct level. Then if firms get better at identifying incompetents, average earnings losses following unemployment will tend to rise. Does this correspond to an increase in insecurity? The argument here is much less clear-cut. The competent will feel more secure because they are less likely to be unfairly labelled as incompetent and suffer an earnings loss. Overall the outcome is hard to judge. However, this argument would seem to imply that the unemployed should be systematically falling in quality relative to the employed as the firms get better at identifying the incompetent and retaining the competent. In fact, there is no evidence, at least on the basis of observables, that this is happening. For example, the proportion of high education (UK A levels+) relative to low education (no qualifications) among the employed has actually fallen relative to that among the unemployed from the late 1970s to the early 1990s in the UK. (see Nickell and Bell, 1996, Table 1). So this explanation is perhaps less persuasive than the others.

Overall, therefore, we feel that in the light of the above discussion it is not unreasonable to suppose that a *ceteris paribus* rise in the earnings losses due to unemployment reflects a rise in general insecurity at work.

The data

The earnings and unemployment data are taken from the UK New Earnings Survey (NES) which has been merged with information from the Joint Unemployment and Vacancy Operating System (JUVOS). The NES is a large sample survey of employees in employment. The sampling frame is based on all individuals whose National Insurance (NI) number ends in a given pair of digits (14). Since NI numbers are issued to every individual prior to starting work and are retained for life, there is a large panel element in the data.

Complete data on earnings are provided for each individual and cover a specific week in April for each year. These data are provided by employers who are legally bound to comply. The data cover hourly and weekly earnings plus detailed information on hours, overtime, age, occupation, industry, region and whether or not the individual was in the same

⁸ Evidence of systematic labour market discrimination against the unemployed is not readily available, although we know from survey evidence that around half of all employers regard unemployment as an undesirable attribute, *per se*, and that the long-term unemployed are systematically disfavoured by, for example, not being selected for interview irrespective of their other characteristics (see Meager and Metcalf, 1987).

job as in the previous year. (Note, s/he can be in a different job with the same employer.) The measure of wages which we use throughout is the weekly pay of those whose pay is unaffected by absence excluding overtime pay divided by weekly hours excluding overtime hours. The idea is to obtain a measure of hourly pay that excludes the overtime element in order to try and eliminate that part of pay which is explicitly sensitive to the business cycle. The alternative is simply to use weekly pay divided by weekly hours but because of the overtime premium, this will vary with hours worked even when the pay schedule is unchanged. Such variation would make the results harder to interpret.

The data on earnings are only available if the individual is in employment on the relevant date and his employer located. Note, that while it is possible and relatively commonplace for an individual's earnings to be unavailable for a given year because he does not have any or his employer is not located, it is very difficult for an individual to disappear completely from the sample unless he dies, emigrates or exits permanently from the labour force because of disability or prison, for example. Merged into these earnings data are the administrative records on unemployment benefits claims (from JUVOS) throughout the previous year. We divide the occupational data into four skill groups along the lines suggested by Elias (1995), the details being provided in Table 2.

Empirical strategy

Taking our panel data set, we divide it into three periods 1982-86, 1987-91, 1992-97 and analyse the following equation,

$$w_{it} = a_i + a_t - \sum_{j=1}^4 b_j D_{ijt} - b D_{it} + \sum_k X_{ikt} a_k + e_{it} \quad (2)$$

Where a_i = individual dummy, a_t = time dummy, w_{it} is ln (hourly earnings), $D_{ijt} = 1$ if the individual completed his first unemployment spell in the sample period up to 3 months ago ($j = 1$), 4-6 months ago ($j = 2$), 7-9 months ago ($j = 3$), 10-12 months ago ($j = 4$); zero otherwise. $D_{it} = 1$ if the first unemployment spell was completed more than 12 months ago; zero otherwise. The X variables include job tenure, age dummies, region dummies. In practice we also include the consequences of a second spell of unemployment with the same structure as the first but, after investigation, we do not include the third or subsequent spells because the numbers are too small to obtain useful results. Note, that because we include time dummies, the earnings loss can be taken to refer to real earnings.

Given the sampling frame individuals exit and enter the sample quite frequently so, despite using a fixed effects estimator, we decide to make further efforts to deal with potential sample selection problems. So, for each year we run a year specific probit explaining the availability of earnings data and then use this to construct Heckman's δ for each sample member for each year. This is included in the estimated equation. The variables included in the probit are age dummies, skill level at the onset of the period and cumulated spells of unemployment. Finally, the equation standard errors are corrected for heteroscedasticity.

Our ultimate aim is to see whether there is any systematic tendency for the β_j and β parameters in Equation (2) to become larger in the later periods. Because the sample is so large ($N \approx 70 K$), we should be able to generate precise estimates of the parameters and to pick up relatively small changes. Furthermore, we can divide the sample by age and skill groups to see if these exhibit any significant changes.

Results

The average earnings losses for men due to unemployment, as generated by the estimated version of Equation (2), are reported in Table 3. The overall pattern is familiar from the results reported in Gregory and Jukes (1997). After the first unemployment spell within the sample period, we see an immediate loss in hourly earnings of somewhere between 10 and 20 percent which is sustained throughout the first year, although there is some tendency for the loss to diminish towards the end of the first year back at work. However, the permanent losses remain quite substantial although it should be borne in mind that since each sample period is only five years long, the “permanent” effect is, in fact, some sort of average of the earnings loss during the period between one and four years after the end of the unemployment spell.

Because the sample size is so large, we have quite precise estimates of the earnings losses and if we make comparisons across periods, we see that the temporary losses in the last period are, on average, just over one-third higher than those in the first period, this being a gap of about 6 standard errors. The permanent loss is about 70 percent bigger in the last period relative to the first, which represents some 10 standard errors. On this basis there seems no doubt that there has been a significant rise in the average earnings losses due to unemployment from the early 80s to the early 90s. The additional losses arising from the second spell are much smaller than the first spell losses but overall they exhibit the same temporal pattern with at least a 50 percent increase from the first period to last period. In nearly all cases most of the jump occurs from the first to the second period remaining more stable thereafter.

There is no obvious macroeconomic reason for this pattern since average claimant count percent unemployment in the three periods is 12.8, 9.7, 11.5 respectively, and unemployment duration exhibits no secular trend. In any event, the costs of becoming unemployed in terms of earnings losses seem to have risen quite sharply over the last two decades. Two other features of these results are worth noting. First, the earnings losses are computed from an equation holding job tenure constant. So these earnings losses are relative to those of an individual with the same fixed effect, the same time effect and the same value of all the other variables including job tenure, ie starting a new job at the same time but without the intervention of a spell of unemployment. If we drop the tenure variables, the earnings losses are very similar, having exactly the same overall pattern. Second, it is striking that the pattern of temporary losses exhibits a slight increase over the first nine months before starting to decline. This suggests that relative to individuals, who do not experience an unemployment spell, the rise in earnings is somewhat slower for the first nine months leading to an increase in losses over this period.

In Table 4, we repeat the exercise but in this case we allow the losses to be influenced by spell duration. Again we see the same picture. Not surprisingly, longer duration unemployment spells are associated with significantly greater earnings losses but, for both short and long unemployment spells, the losses in the 1992-97 period tend to be around one-third higher than those in the 1982-86 period. In Table 5, we look at how earnings losses vary with age dividing the sample into three age groups, using the age at the beginning of each period. The most notable feature of the results is that there is no evidence of a rise in earnings losses for the young which are, not unexpectedly, much smaller than the earnings losses for older workers in any event. By contrast, the rise in earnings losses for prime age men is really substantial being of the order of 70 percent with a somewhat smaller rise in earnings losses for older men.

When we divide up the sample into the four skill levels at the beginning of each period, we find a number of results worth noting in Table 6. First, the earnings losses tend to

be higher, the higher the skill level which is consistent with a specific human capital explanation. Second, the increase in earnings losses as we move into the later periods is more obvious for the two higher skill groups than for the two lower skill groups. Concentrating on the permanent effects, the bottom skill group (level 1) exhibits no systematic increase in income loss at all as well as being relatively small. The next level skill group (level 2) also exhibits little secular increase over the three periods, although overall the losses are somewhat larger than in the bottom group. By contrast, the top two skill groups suffer wage losses due to unemployment which are substantially larger in the last period than in the first. The temporary losses are over 40 percent larger in the last period than in the first and the permanent losses tend to be around 80 percent bigger. So in the final period, the permanent losses suffered by the top skill group correspond to a wage fall in excess of 25 percent. In the light of these facts, it is worth reporting that the percentage of each skill group suffering one or more unemployment spells in the three periods are on average around 17, 16, 15 and 11 going from skill level 1 up to skill level 4. Furthermore there is no significant tendency for this to increase over time which is consistent with the analysis in the previous section.

Summary

While in the previous section we found no serious evidence of any systematic increase in the chances of becoming unemployed, we find here that there has been a strong tendency for the costs of unemployment to increase, particularly for those in the older age groups and the higher skill groups. The losses in hourly earnings consequent on unemployment for men outside the bottom skill group and the youngest age group have risen by 30 percent or more from the early 1980s to the early 1990s with the largest losses affecting the highest skill group. In the next section we turn to the pattern of earnings for those who remain in the same job.

4. How Common Are Falls in Real Wages?

Individual real wages are subject to substantial fluctuations, year on year, even for people who are employed continuously in the same job. For example, in a typical year, at least 7 percent of continuously employed people will experience a real wage fall of more than 10 percent, and around 40 percent of these will experience a decline of over 30 percent. How, then, does the probability of a substantial decline in real wages relate to insecurity? If fluctuations in real earning are known in advance, they certainly make a lesser contribution to insecurity and presumably no contribution at all if capital markets are perfect⁹. However, it seems reasonable to suppose that most year-on-year fluctuations in real wages are not accurately known in advance. Consequently, an increase in the chances of a substantial fall in real wages will enhance feelings of insecurity, particularly in the presence of commonplace capital market imperfections.

So what has happened to real wage fluctuations in recent years? In Table 7, we record the percentage of men who have faced one year falls in real hourly pay exceeding 10 percent and 30 percent respectively where real pay is generated by normalising on the retail price index. We consider three groups: those who were continuously employed in the same job, those who changed jobs without an intervening spell of unemployment and those with an

⁹ Whether someone facing known fluctuations in income when capital markets are imperfect feels more insecure than someone whose income remains constant is a moot point. In any event, someone facing uncertain fluctuations in income certainly feels more insecure under any commonness view of insecurity.

intervening spell of unemployment¹⁰. Of course there is another category whose wage in the second period is not available because they are still jobless, having been employed in the first period. As we have already seen, they will generally have suffered a substantial and increasing earnings loss on returning to work.

Turning first to those continuously employed in the same job, we see from the first two columns of Table 6 that there has been a slow but steady rise in the percentage of men facing substantial falls in real wages. In the period 1982-86, 7.6 (3.0) percent of men suffered a fall in real hourly pay of 10 (30) percent and this rose in 1992-96 to 9.5 (4.3), a highly significant increase. For those who changed jobs without an intervening spell of unemployment, the chances of a substantial fall in real pay are somewhat larger, despite the fact that the median change for this group tends to be around 4 percentage points higher than for the continuously employed. Nevertheless, the temporal pattern is much the same with a steady increase in the chance of a 10 (30) percent fall from 12.0 (6.4) percent to 14.4 (7.9) percent. Finally, and not surprisingly in the light of the previous section, we see that the chances of a large fall in real pay are very high for those who shift jobs with an intervening spell of unemployment. The average probability of a 10 (30) percent fall in real wages for men was 33.5 (21.0) percent in 1982-86 rising to 40.4 (27.9) percent in 1992-96.

This consistent increase in the chances of a substantial fall in real wages raises a number of questions. First, we know that over the sample period there has been a substantial increase in the dispersion of the level of earnings. So is this related to our findings? In fact, there is no particular reason why it should be. The distribution of year-on-year changes in real earnings bears no specific relation to the levels distribution, which can widen even if the distribution of changes is becoming more compressed. Second, does the increase in the chances of a substantial fall in real wages in fact reflect an increase in the dispersion of changes? In Table 8, we see that the answer is no, for the proportion of changes within the " 10 or " 30 percent range exhibits no systematic decline and so there has been no secular increase in dispersion.

The third question concerning the systematic rise in the chances of a substantial fall in real wages is whether or not these falls simply reflect a transitory change that will shortly be reversed. To pursue this issue we repeat the exercise of Table 7 using 3-year averages rather than year-on-year changes. The idea here is as follows. For those with no job change, the change in period t is the difference between a 3-year average of real hourly pay after t and the same before t , for all those in the same job for six consecutive years. For those with a job change, we take all those who change jobs at t and then take the difference between the three year averages after and before t . The results are reported in Table 9 and exhibit much the same pattern as those in Table 7. Of course, the numbers are smaller and those in the last two columns are very unreliable because of the small sample. Nevertheless, by the 1990s, over 5 percent of individuals continuously employed in the same job could expect to face a long-run fall in real hourly pay of more than 10 percent, the numbers having risen by more than 2 percentage points since the mid 1980s. So even if we focus on long-run falls in real pay, we find a significant increase in incidence over the relevant period.

The final question worth asking is whether or not this increase relates to changes in aggregate factors. We know that it does not relate to the cycle since the average unemployment rates over the three periods exhibit no secular pattern (they are 12.8, 9.7, 11.5 respectively). However, more plausible is the possibility that the increase relates to the aggregate movements in real pay. If real pay increases are lower in the later periods, this might explain why the chances of a real pay cut are larger. To pursue this, we simply repeat

¹⁰ It is worth recalling that a job change does not necessarily mean a change of employer since a significant change of post within the firm is counted as a job change.

the analysis reported in Tables 7 and 9 but first we subtract off the median rise in real hourly pay each year from all the observations¹¹. The consequences are reported in Table 7a and Table 9a and reveal immediately that this explains most, and sometimes all, of the change. So the picture we have is one where there has been no secular trend in the dispersion of changes but, in each period, the distribution of changes moves to the left or right as the median moves and the proportion of individuals subject to a substantial real pay cut goes up and down with these changes in the median. This does not, of course, affect the fact that the chances of a substantial real pay cut have risen over the last fifteen years but it does suggest this is not part of a very long-run trend but merely reflects the fact that real pay rises happen to be smaller than they were in the mid-1980s.

In Table 10, we break down the chances of a 10 percent fall in real pay for men continuously employed in the same job by age group. A clear pattern emerges. Leaving aside teenagers, in the early period (1982-86) the over 50s tended to have somewhat lower chances than the under 50s of a fall in real wages of more than 10 percent. By 1992-96 the over 50s have a higher chance of such a fall, although this probability rose significantly for all age groups. So there is some evidence that older men have been harder hit by an increase in this type of earnings insecurity than younger men. Because the sample sizes are much smaller, we do not repeat this exercise for those who change jobs. This pattern carries over to long-term falls in real pay (using 3-year averages) with the chances of a 10 percent decline rising by about 100 percent (from 3.3 to 6.4) for the over 60s, compared to a rise of less than 60 percent (from 2.4 to 3.7) for the under 30s. If we now repeat the exercise subtracting the median increase each year from every observation, we find that the rise in the chances of a large real pay cut for the young is completely explained by the overall decline in a real pay rise over the period. This does not apply to the old for whom the “corrected” chances of a 10 percent real pay cut still rises by 2 percentage points from the start to finish.

Perhaps the most interesting breakdown of these data is generated by looking at skill levels in Table 11. What we see is first that the bottom skill group always has the highest, and the top skill group always has the lowest, chance of a fall in real pay in excess of 10 percent. This is despite the fact that their respective pay levels are already relatively low and relatively high in the first place. Second, the chances of a 10 percent fall in real pay has gone up significantly for all groups over the relevant period. Third, and most interesting, we see that the top skill group has had the biggest increase in the probability of a 10 percent fall in real pay and is catching up with the other groups. So, the relative position of the top skill group has worsened since the early 1980s in this regard, just as the top skill group has come out relatively badly when it comes to the increasing earnings losses due to unemployment. This is reinforced if we consider long-run real pay movements (3-year averages) where again the high skill groups lose out more than the low skill groups in terms of their increase in the chances of a large real pay cut. Then if we subtract off the overall median real pay increase from each observation and repeat the exercise, for the low skill groups there is nothing left – the chances of a ten percent fall in the median corrected real pay have not risen at all. By contrast, for the highest skill group their remains a significant increase of nearly 3 percentage points. This means that variations in the overall rise in median pay explain only a small proportion of the increase in the chance of a ten percent fall in real pay for this group. (Note in Table 11 the high skill increase is around 3.5 percentage points.)

¹¹ We use the median rather than the mean because there are some very extreme real wage changes, particularly increases, which we feel tend to distort the measure of the simple average change.

Summary

For men who are continuously employed and for men who change jobs, we see a clear and significant increase in the chances of a substantial year-on-year decline (10 percent or more) in real hourly wages over the period from the early 1980s until the mid 1990s. Over this period, the probability of a 10 percent real wage decline has risen by around 30 percent for continuously employed men and by around 20 percent job changers. These changes are not due to an overall increase in the dispersion of wage changes and they show up just as strongly if we consider “long-run” changes (3-year averages). However, a good part of this overall change is due to the decline in the median rate of real pay rises over the relevant period. Older workers and the top skill group have seen a worsening of their relative position in this regard with those men in the top skill group who are continuously employed seeing a 60 percent rise in their chances of a 10 percent fall in real wages, year-on-year. Furthermore, this particular rise is not due to the overall fall in the median rate of real pay rises. Despite this, men in the lower skill groups remain more insecure and in the most recent period men continuously employed in the same job in the bottom skill group are still nearly 50 percent more likely to experience a 10 percent year-on-year drop in real hourly pay than similar men in the top skill group.

5. Conclusions

We have looked at three aspects of the job insecurity facing British men in the last two decades. The probability of becoming unemployed, the cost of unemployment in terms of real wage falls and the probability that the continuously employed will experience substantial real wage declines. The following facts emerge:

- i) There is little or no evidence of any trend increase in the chances of men becoming unemployed over the last twenty years, either on average or within any particular occupational or skill group.
- ii) There has been a strong tendency for the costs of unemployment in terms of wage losses to increase for all men except those in the lowest skill group. The losses in hourly earnings consequent on unemployment for men outside the bottom skill group have risen by around 40 percent or more from the early 1980s to the mid-1990s, with the largest losses affecting the highest skill group.
- iii) For men who are continuously employed in the same job and for those who change jobs, the chances of a substantial year-on-year decline (10 percent or more) in real hourly pay have increased by 20 to 30 percent from the early 1980s to the mid 1990s. Older workers and those in the top skill group have seen a worsening of their relative position in this regard. Despite this, those in the bottom skill group are still far more likely to experience a substantial year-on-year drop in real wages than those in the top skill group. A similar pattern holds for longer-term (3-year average) changes in real hourly pay.
- iv) The overall changes in iii) are not due to an increase in the dispersion of wage changes but to the fact that the median annual rise in real pay in the 1990s is smaller than that in the mid-1980s. This suffices to shift the whole distribution of wage

changes to the left, thereby increasing the chances of a 10 percent decline. However this does not explain the changes for the old and the high skilled.

Overall, therefore, there has been a rise in job insecurity for British men since the early 1980s. This has come about not because of a rise in the chances of losing their jobs but because the cost of job loss has risen and, for the continuously employed, the probability of a substantial year-on-year fall in real wages has gone up. Finally, insecurity has gone up by more for those in the highest skill groups.

Table 1
Explaining the Percentage Probability of an Employee
being Unemployed Twelve Months Later
Dependent Variable: p_{it}

Constant	-1.81
u_t	0.38 (6.7)
SEG 2	-0.17 (0.1)
SEG 3	0.67 (0.6)
SEG 4	6.55 (5.6)
SEG 5	3.07 (2.6)
SEG 6	3.88 (3.3)
SEG 7	7.27 (6.2)
(SEG 1) t	0.034 (0.6)
(SEG 2) t	0.064 (1.1)
(SEG 3) t	0.087 (1.5)
(SEG 4) t	-0.030 (0.5)
(SEG 5) t	-0.030 (0.5)
(SEG 6) t	0.009 (0.2)
(SEG 7) t	-0.006 (0.1)
NT	126
\bar{R}^2	0.18

Notes:

p_{it} is the proportion of men working in SEG i in $t-1$ who are unemployed twelve months later.

SEG i is a dummy variable taking the value one if they were working in SEG i , zero otherwise.

(SEG i) t is the interaction between SEG i and a time trend.

SEG1 = Employers, managers, professionals

SEG2 = Intermediate non-manual

SEG3 = Junior non-manual

SEG4 = Personal service workers

SEG5 = Foreman, supervisors, skilled manual

SEG6 = Semi-skilled manual

SEG7 = Unskilled manual

The unit of observation is the SEG for the years 79, 81, 83, 84-98, so there are $7 \times 18 = 126$ observations. The regression is estimated by OLS with SEG dummies (ie a fixed effects model) and the absolute t ratios are in parentheses.

Table 2
Skill levels based on the Standard Occupational Classification

Skill Level	Major Groups	Constituent Minor Groups (2 digit)
Level 4	Managers and administrators (excluding office managers and managers / proprietors in agriculture and services). Professional occupations	10, 11, 12, 15, 19, 20-27, 29
Level 3	Office managers and managers / proprietors in agriculture and services Associated professional and technical occupations Craft and related occupations Buyers, brokers, sales reps	13, 14, 16, 17 30-39 50-59 70,71
Level 2	Clerical, secretarial occupations Personal and protective service occupations Sales occupations (except buyers, brokers, sales reps) Plant and machine operatives Other occupations in agriculture, forestry, fishing	40-46, 49 60-67, 69 72, 73, 79 80-89 90
Level 1	Other elementary occupations	91-95, 99

Source: Elias (1995)

Table 3
The Impact of Unemployment on the Hourly Pay of Men
Hourly Earnings Loss (%)

	1982-86	1987-91	1992-97
Impact of 1 st unemployment spell			
After 1-3 months	12.6 (0.98)	17.9 (1.23)	14.7 (1.75)
4-6 months	13.2 (0.77)	20.0 (0.87)	18.4 (0.96)
7-9 months	13.8 (0.73)	20.4 (0.81)	18.7 (0.82)
10-12 months	11.3 (0.72)	17.3 (0.79)	17.3 (0.80)
Permanent	9.0 (0.50)	14.4 (0.63)	15.5 (0.68)
Impact of 2 nd unemployment spell			
After 1-3 months	0.4 (1.5)	5.8 (1.57)	7.4 (2.01)
4-6 months	3.4 (1.14)	9.4 (1.30)	7.2 (1.07)
7-9 months	5.0 (1.08)	8.5 (1.17)	7.8 (0.89)
10-12 months	2.3 (1.08)	5.9 (1.16)	6.3 (0.93)
Permanent	1.8 (0.74)	5.4 (0.88)	5.2 (0.67)
NT	387571	386472	416111
R ²	0.199	0.167	0.115

Notes:

- ii) These results are based on regressions which include fixed individual effects (ie the regressions are within groups), time dummies, region dummies, age dummies, job tenure, job tenure², Heckman's δ .
- iii) Standard errors in parentheses.
- iv) Only unemployment spells in excess of ten days are counted.

Table 4
The Impact of Unemployment Spells on the Hourly Pay of Men
Analysis of Spell Duration
Hourly Earnings Loss (%)

	1982-86	1987-91	1992-97
Impact of 1st Spell of Unemployment			
After 1-3 months	11.7 (1.13)	17.0 (1.40)	9.1 (2.05)
4-6 months	11.8 (0.89)	19.3 (0.99)	16.6 (1.01)
7-9 months	12.8 (0.86)	19.9 (0.90)	16.1 (0.94)
10-12 months	9.9 (0.84)	16.1 (0.89)	14.3 (0.90)
Permanent	7.2 (0.54)	12.8 (0.66)	13.0 (0.64)
Additional Earnings Loss if Spell Exceeds 6 Months (%)			
After 1-3 months	4.5 (2.14)	9.0 (2.92)	17.7 (3.61)
4-6 months	7.1 (1.67)	8.4 (2.16)	7.3 (1.91)
7-9 months	5.5 (1.49)	8.0 (2.05)	9.8 (1.53)
10-12 months	6.7 (1.45)	10.3 (2.03)	10.9 (1.56)
Permanent	8.3 (1.05)	11.4 (1.79)	9.9 (1.00)

Notes: See Table 3, notes i), ii), iii)

Table 5
The Impact of Unemployment Spells on the Hourly Pay of Men
Analysis by Age
Hourly Earnings Loss (%)

	1982-86	1987-91	1992-97
Age ≤ 30			
Impact of 1 st Spell of Unemployment			
After 1-3 months	11.0 (1.4)	12.8 (1.6)	5.6 (2.2)
4-6 months	10.3 (1.1)	13.2 (1.2)	7.5 (1.3)
7-9 months	11.8 (1.0)	12.5 (1.1)	9.1 (1.0)
10-12 months	7.3 (1.0)	9.6 (1.1)	4.8 (1.0)
Permanent	6.3 (0.8)	7.7 (0.9)	6.7 (0.8)
31 ≤ Age ≤ 50			
Impact of 1 st Spell of Unemployment			
After 1-3 months	16.1 (1.6)	19.7 (2.1)	20.7 (3.1)
4-6 months	17.3 (1.3)	22.7 (1.4)	27.5 (1.5)
7-9 months	14.7 (1.1)	24.8 (1.3)	22.0 (1.3)
10-12 months	13.7 (1.1)	20.7 (1.2)	25.0 (1.3)
Permanent	11.4 (0.8)	17.7 (1.0)	20.9 (1.0)
51 ≤ Age			
Impact of 1 st Spell of Unemployment			
After 1-3 months	22.6 (2.5)	19.5 (3.4)	30.4 (6.5)
4-6 months	22.5 (2.7)	28.6 (2.5)	22.6 (3.5)
7-9 months	21.8 (2.0)	28.1 (2.6)	27.6 (3.6)
10-12 months	23.3 (2.3)	27.5 (2.2)	33.9 (3.7)
Permanent	18.5 (1.4)	25.3 (1.8)	28.5 (2.4)

Notes: See Table 3, notes i), ii), iii)

Table 6
The Impact of Unemployment Spells on the Hourly Pay of Men
Analysis by Skill Level
Hourly Earnings Loss (%)

	1982-86	1987-91	1992-97
Skill Level 1 (low skill)			
Impact of 1 st Spell of Unemployment			
After 1-3 months	13.6 (4.9)	7.4 (4.4)	11.0 (8.5)
4-6 months	7.2 (3.4)	17.7 (2.9)	16.5 (5.2)
7-9 months	13.7 (3.4)	13.4 (2.7)	1.1 (3.6)
10-12 months	11.0 (3.3)	11.9 (2.7)	13.4 (4.6)
Permanent	6.3 (2.2)	11.4 (2.1)	4.1 (2.4)
Skill Level 2 (low intermediate)			
Impact of 1 st Spell of Unemployment			
After 1-3 months	11.0 (2.7)	17.2 (2.0)	14.9 (4.1)
4-6 months	11.5 (1.5)	18.5 (1.4)	13.4 (2.1)
7-9 months	12.0 (1.5)	17.2 (1.3)	10.7 (1.9)
10-12 months	12.1 (1.5)	15.9 (1.2)	10.8 (1.9)
Permanent	9.0 (0.9)	12.3 (0.9)	10.6 (1.1)
Skill Level 3 (high intermediate)			
Impact of 1 st Spell of Unemployment			
After 1-3 months	8.8 (2.4)	19.1 (2.2)	12.1 (3.6)
4-6 months	11.7 (1.7)	19.2 (1.6)	17.9 (2.0)
7-9 months	10.5 (1.5)	20.3 (1.5)	15.8 (2.1)
10-12 months	8.9 (1.4)	17.3 (1.5)	17.0 (2.1)
Permanent	8.8 (0.9)	15.0 (1.1)	16.8 (1.2)
Skill Level 4 (high skill)			
Impact of 1 st Spell of Unemployment			
After 1-3 months	18.2 (5.5)	24.6 (4.6)	25.9 (7.2)
4-6 months	20.9 (3.7)	20.7 (2.8)	31.7 (4.5)
7-9 months	17.8 (3.0)	25.6 (2.6)	25.0 (3.4)
10-12 months	22.8 (3.9)	21.8 (2.7)	27.7 (3.4)
Permanent	15.0 (1.7)	20.7 (2.0)	26.7 (2.1)

Notes: See Table 3, notes i), ii), iii) –
iv) for precise definitions of skill levels, see main text and Table 2.

Table 7
Percentage of Men Facing Significant One Year Falls in Real Hourly Pay

	No Job Change		Job Change, No Unemployment		Job Change, Unemployment	
	Down > 10%	Down > 30%	Down > 10%	Down > 30%	Down > 10%	Down > 30%
1982	7.6	3.1	12.2	6.2	37.6	22
1983	7.2	3.1	10.4	5.1	30.4	19.8
1984	9.1	3.5	14.1	8	33.8	20.3
1985	6.5	2.6	10.7	5.6	31.7	19.1
1986	7.4	3	12.6	7	33.6	23.2
1987	7	3	12.1	6.2	32.1	20.4
1988	9.4	4.1	13.6	7.5	34.9	25
1989	10.3	4.4	14.7	8.2	42.8	28.9
1990	9	4.1	12.1	6.8	40	26.6
1991	8.2	3.6	12.2	6.5	37.2	25.6
1992	7.4	3.4	11.8	6.6	43.5	31.8
1993	9.5	4.1	14.6	7.2	40.2	25.5
1994	11	4.9	14.9	8.1	43.9	31.4
1995	10.5	4.9	15	8.4	35.6	24.8
1996	9.2	4.4	15	9	37.3	24.3
Approx se (%)	0.125	0.11	0.48	0.37	2.14	1.8
Averages						
1982-86	7.6	3	12	6.4	33.5	21
1987-91	8.8	3.8	13	7.2	37.1	25.1
1992-96	9.5	4.3	14.4	7.9	40.4	27.9
Average Sample Sizes						
1982-86	51299		3924		614	
1987-91	50298		6320		571	
1992-96	48502		4535		387	

Source: UK New Earnings Survey. Note: The standard error (se) is based on the formula $(\hat{p}(1 - \hat{p}) / n)^{1/2}$ where \hat{p} is the proportion and n is the sample size. It refers to the proportion for a single year. The change in real hourly pay refers to changes for a given week in April from one year to the next.

Table 7a
Percentage of Men Facing a 10 Percent Real Pay Cut
After Controlling for the Median Pay Rise

	No Job Change	Job Change No Unemployment	Job Change Unemployment
Averages			
1982-86	9.6	22.5	34.8
1987-91	10.5	22.4	35.1
1992-96	10.5	20.5	35.8
Approx. se (%)	0.13	0.5	2.1

Note: In this table, we have taken the data summarised in Table 7 and for each year we have subtracted off the median real pay increase from every observation and then computed the probability of a 10 percent real pay reduction.

Table 8
Indicators of Dispersion for Those with no Job Change
Percentage of men with change in real pay (^aRP)
Bounded by " 10% and " 30%

	- 10% < ARP < 10%	- 30% < ARP < 30%
Averages		
1982-1986	70.3	86.5
1987-1991	68.7	85.4
1992-1996	71.2	85.7

Table 9
Percentage of Men Facing “Long-run” Falls in Real Hourly Earnings

No Job Change		Job Change, No Unemployment		Job Change Unemployment		
	Down > 10%	Down > 30%	Down > 10%	Down > 30%	Down > 10%	Down > 30%
1984	3.3	1	5.6	2	33.3	22.2
1985	2.9	0.9	4.26	2.3	24.5	17
1986	2.8	0.9	5.42	2.6	27.2	18.5
1987	3.2	0.8	5.5	2.1	29.6	20.4
1988	4.9	1.2	5.8	3	45.7	28.3
1989	5.8	1.6	7.2	2.9	47.6	28.6
1990	5.7	1.7	5.8	3.2	43.6	20
1991	5.2	1.6	7.2	3.3	26	16
1992	4.7	1.6	7.3	4	50.8	32.3
1993	5.6	1.7	8.8	4.2	40.4	25
1994	6.3	2	9.5	5.1	67.5	47.8
Approx se (%)	0.16	0.09	0.71	0.55	6.5	5.9
Averages						
1984-87	3.1	0.9	5.1	2.2	28.3	19.3
1988-91	5.4	1.5	6.5	3.1	40.4	22.8
1992-94	5.5	1.8	8.6	4.5	52.2	34.4
Average Sample Sizes						
1982-85	16438		873		58	
1986-89	15516		1271		65	
1990-92	17666		893		54	

Table 9a
Percentage of Men Facing a “Long-run” 10 Percent Real Pay Cut
After Controlling for the Median Pay Rise

	No Job Change	Job Change No Unemployment	Job Change Unemployment
Averages			
1982-85	10.5	27.8	35.7
1986-89	13.1	24.9	31.1
1990-98	11.3	21.5	31.3
Approx. se (%)	0.16	0.71	6.5

Note: In this table, we have taken the data summarised in Table 7 and for each year we have subtracted the median real pay increase from every observation before proceeding.

Table 10
Percentage of Men Facing One Year Falls in Real Hourly Pay of
10 Percent or More No Job Change: Analysis by Age

Age	16-20	21-30	31-40	41-50	51-60	60+
Average						
1982-86	5.3	7.6	8	8.2	7.4	7.4
1987-91	5.9	8.8	9.2	9.3	8.6	8.8
1992-96	6.6	9.1	9.7	9.9	10.1	10.2
Approx se (%)	1.4	0.27	0.26	0.27	0.32	0.83

Source: UK New Earnings Survey Note: the standard error (se) is based on the formula $(\hat{p}(1 - \hat{p}) / n)^{1/2}$ where \hat{p} is the proportion and n is the sample size. It refers to the proportion for a single year. Obviously the average over 5 years is more precise. The change in real hourly pay refers to changes for a given week in April from one year to the next.

Table 11
Percentage of Men Facing One Year Falls in Real Hourly Pay
of 10 Percent or More Analysis by Skill Group

Skill	Level 1	Level 2	Level 3	Level 4
Group	(low skill)	(low intermediate)	(high intermediate)	(high skill)
No Job Change				
Average				
1982-86	10.6	8.1	7.4	5.2
1987-91	12.2	8.7	8.9	7.3
1992-96	12.5	9.3	9.7	8.8
Approx se (%)	0.51	0.21	0.22	0.32
Job Change, No Unemployment				
Average				
1982-86	14.7	11.7	12.6	10.6
1987-91	16	13.3	12.8	12.1
1992-96	18.4	13.3	14.7	14.4
Approx se (%)	2.15	0.79	0.82	1.1

Source: UK New Earnings Survey. Notes: the standard error (se) is based on the formula $(\hat{p}(1 - \hat{p}) / n)^{1/2}$ where \hat{p} is the proportion and n is the sample size. It refers to a single year. The change in real hourly pay refers to changes for a given week in April from one year to the next. For precise definitions of the skill levels, see the main text.

References

- Addison, J.T. and Portugal, P. (1989), 'Job Displacement, Relative Wage Changes and Duration of Unemployment', Journal of Labor Economics, 7 July, pp.281-302.
- Burgess, S. and Rees, H. (1996), 'Job Tenure in Britain 1975-92', Economic Journal, 106, March, 334.44.
- Chowdhury, G. and Nickell, S. (1985), 'Hourly Earnings in the United States: Another Look at Unionization, Schooling, Sickness, and Unemployment using PSID Data', Journal of Labor Economics, 3 (1), Part I, January, pp.38-69.
- Elias, P. (1995), 'Social Class and the Standard Occupational Classifications' in D. Rose (ed.) A report on Phase 1 of the ESRC review of the OPCS Social Classification. Economic and Social Research Council: Swindon.
- Green, F., Felstead, A. and Burchall, B. (1998), 'Job Insecurity and the Difficulty of Regaining Employment: An Empirical Study of Unemployment Expectations', Leeds University Business School, mimeo.
- Gregg, P. Knight, G. and Wadsworth, J. (1997), 'Heaven Knows I'm Miserable Now: Job Insecurity in the British Labour Market', Centre for Economic Performance, Working Paper No. 892, London School of Economics.
- Gregg, P. and Wadsworth, J. (1996), 'Mind the Gap, Please? The Changing Nature of Entry Jobs in Britain', Centre for Economic Performance, Working Paper No. 796, London School of Economics.
- Gregory, M. and Jukes, R. (1997), 'The Effects of Unemployment on Subsequent Earnings: A Study of British Men 1984-94', Institute of Economics and Statistics Leverhulme Programme on the Labour Market Consequences of Technical and Structural Change, DP No. 21, Oxford University.
- Jacobsen, L., Lalonde, R. and Sullivan, D. (1993), 'Earnings Losses of Displaced Workers', American Economic Review, 83, September, pp.685-709.
- Meager, N. and Metcalf, H. (1987), 'Recruitment of the Long Term Unemployed', Institute of Manpower Studies, IMS Report No. 138.
- Nickell, S. and Bell, B. (1996), 'Changes in Distribution of Wages and Unemployment in OECD Countries', American Economic Review, 86, pp.302-308.
- Nickell, S. and van Ours, J. (1999), 'The Netherlands and the United Kingdom: A European Unemployment Miracle?', forthcoming in Economic Policy.
- OECD (1994), *Jobs Study*, OECD: Paris.
- Ruhm, C.J. (1991), 'Are Workers Permanently Scarred by Job Displacements?', American Economic Review, 81, March, pp.319-324.

Swain, P. and Podgursky, M. (1991), 'The Distribution of Economic Losses Among Displaced Workers: A Replication', Journal of Human Resources, 26, Fall, pp.742-755.

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