association with unemployment. Since the greatest employment effects should be felt by youth, the conclusion regarding employment rates is telling:

> It is important to note that these estimated effects are relatively insignificant in terms of explaining the large decline that has occurred in the teenage employment-population ratio in some countries . . . the substantial difference across countries in teenage employment trends can only be marginally attributed to differences in the evolution of minimum wages . . . (OECD, 1998, p. 48).

If wage rigidity were as important as claimed by the conventional wisdom, unemployment and its change over time ought to be strongly negatively associated with earnings inequality, unit labor costs, and wage shares (in total income). In fact, as discussed above (Figures 7.3 and 7.4), the data shows no clear tradeoff between changes in unemployment and changes in earnings inequality. Nor is there a negative association for levels. For example, for 17 countries, the simple correlation coefficient between a standard measure of earnings inequality (D9/D1) and unemployment rates for the early 1990s was .101, a statistically insignificant relationship. Substituting an alternative inequality measure, the ratio of median earnings to the 10th percentile level (D5/D1), shows an even lower association (.085). Nor have aggregate wage costs been rising: wage shares in nearly all countries have been on a downward trend since the early 1980s (OECD, 1997, figure 10).

The two unemployment insurance variables are key measures of 'employment-unfriendly' labor market institutions. While the replacement rate and benefit duration variables are usually found to be statistically significant in the published literature, their importance and robustness are open to question. Table 7.1 presents simple OLS regression results for different measures of unemployment over the course of the last decade or so. Panel A shows results for 1989–94 for 20 countries. The institutional variables are the standard ones with one exception. The usual duration measure is subjectively defined, from 0 to 4. Spain provides an example of the difficulty of constructing such a measure. The standard Nickell-Layard data set assigns Spain a benefit duration value of 3.5 years. But as Munoz de Bustillo comments,

> The 3.5 years of duration considered is only the maximum, subject to strict eligibility criteria and associated with a much lower benefit replacement ratio. In fact 40% of the unemployed receiving unemployment compensation have benefit duration of 1 year or less. On the other hand 44% qualify for a benefit duration from 1.5 to 2 years . . . (Bustillo, 2001).

This suggests that perhaps a 1.5 value would be more appropriate. Separate results are shown for a duration variable in which Spain's value has been changed from 3.5 to 1.5.

Panel A of Table 7.1 shows a fairly similar pattern to the Fitoussi et. al. results reported above. Union density is insignificant for both total and long-term unemployment. Coordination tends to reduce unemployment, but is insignificant for long-term unemployment. Active labor market policies also show the standard negative effect, but again the effects are not reliably measured (low t-statistics). Typically, the two unemployment insurance benefits (UIB) measures have the expected positive effects on total unemployment. The replacement rate is insignificant for long-term unemployment. A key result for our purposes appears in columns 2 and 5, where Spain's duration measure has been changed from 3.5 to 1.5. The coefficient, 't' statistic, and adjusted R squared all drop, with a substantial decline in the explanatory power of the total and long-term equation (over 10 percentage points in each case). As would be expected, the declines are slightly smaller if a value of 2 is used (not shown). Clearly, the results are highly sensitive to judgments about what duration value to attach to each country.

Panel B shows results for tests of the same equation for unemployment in 1995 and 2001, and for young female unemployment in 1995. For 1995, the standard institutional variables perform poorly. Our alternative duration measure is not significant at the 10 per cent level for 1995, and neither duration measure is statistically significant for young female unemployment in 1995 or for total unemployment in 2001. The predictive power of this key institutional measure declines sharply between the early and late 1990s.

But even in the 1980s and early 1990s, when the benefits duration measure appears to have been most strongly associated with unemployment, and ignoring questions concerning the construction of this variable, the evidence for this lynchpin of the conventional story is less convincing that at first glance. Emphasizing the centrality of this disincentive for job search, Layard et. al. write that 'It is noticeable … that all the countries where long-term unemployment has escalated have unemployment benefits of some kind that are available for a very long period, rather than running out after 6 months (as in the USA) or 14 months (as in Sweden)' (Layard et al., 1994, pp. 59–62) This is certainly one way to view the data. But their figure can also be read to show that the nine countries with 'indefinite benefits' have widely varying propensities for long-term unemployment, with the share of the unemployed out of work for over a year ranging from 20 per cent (Finland) to over 70 per cent (Belgium). Through this lens, the fit between benefits and long-term unemployment in the mid-1980s does not look very close. Then, of course, there is the problem of causation. It would be