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Abstract

In this paper, we provide a set of comparable estimates of aggregate monthly job-finding and separation rates for twenty-seven OECD (Organisation for Economic Co-operation and Development) countries; these estimates can be used for the cross-country calibration of search models of unemployment. We find that cross-country differences in job-finding rates are much greater than those in separation rates. Our results are quantitatively and qualitatively in line with those published in previous studies; however, they cover a much larger set of countries. We combine our estimates with evidence on unemployment and labor force participation rates to impute steady-state worker flows for twenty-three of the countries in our sample.

Key words: labor markets, search theory, worker flows

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1 Introduction

Search models of the labor market, based on Diamond (1982) and Mortensen and Pissarides (1994), have become the workhorse of modern macroeconomic theories of unemployment.[1] The key inputs into these models are the job-finding and separation rates.

As described in Nickell (1997), labor market outcomes differ substantially across OECD countries. Using search models to understand the reasons for cross-country labor market differentials therefore requires consistent estimates of the job-finding and separation rates. In this paper, we provide a set of comparable estimates of aggregate monthly job-finding rates and separation rates for 27 OECD countries that can be used for the cross-country calibration of search models of unemployment.

The aggregate job-finding rate is defined as the ratio of the flow from another activity into employment to the number of people seeking jobs. As Hall (2005a) points out, finding the denominator is not easy because flows from employment to employment and out of the labor force to employment are not that easy to measure. In practice, empirical studies focus on the fraction of unemployed persons that flow out of unemployment as the aggregate job-finding rate and the fraction of workers that leave their jobs as the aggregate separation rate. These are the definitions we use in this paper. At the end of our empirical analysis we discuss the implications of our estimates when one takes into account various more-difficult-to-observe worker flows, like job-to-job transitions and flows into and out of the labor force.

The estimates we provide are obtained by applying the Generalized Method of Moments (GMM) to the implications of the steady-state of a search model of the labor market for the aggregate unemployment duration distribution, as well as the aggregate job tenure distribution. The cross-country comparable data on these distributions that we use are, in most part, taken from OECD (2006a,b). Our sample covers the period 1968-2004.

There are several previous estimates of aggregate job-finding and separation rates that were intended for use in macroeconomic models of unemployment and search. For the U.S., there are estimates by Shimer (2005), Hall (2005a,b), and Fujita and Ramey (2006). Ridder and van den Berg (2003) provide one of the few comparable estimates of job-finding rates across countries. Their

[1] Such theories have recently found applications in the analysis of monetary policy. See Gertler and Trigari (2006) and Krause and Lubik (2007).

sample countries are France, Germany, the Netherlands, the U.K., and the U.S. As for separation rates, Jolivet, Postel-Vinay, and Robin (2006) provide estimates of separation rates for a sample of 11 countries. None of these studies, however, provides a set of estimates for a sample of countries as comprehensive as ours.

Alternatively, many applied microeconomic studies try to estimate the effect of labor market policies, most notably the level and duration of unemployment benefits, on workers' rate of exit out of unemployment.[2] These studies generally focus on estimating the re-employment probabilities around the expiration of jobless benefits, and, in particular, measuring the magnitude of the spike in the exit rate when expiration occurs. They use information on individuals who have taken up unemployment insurance benefits, as opposed to data on the total population of unemployed, which is what we use for our analysis. Moreover, though such studies are crucial for the development and evaluation of (active) labor market policies, they are of less interest for macroeconomic studies of unemployment in which cross-worker heterogeneity is generally ignored, or at least considered of second-order importance.

Since the definitions of different labor market states might vary across data sets, it is not a trivial task to obtain consistent and comparable estimates of job-finding rates for different countries from micro data. Similar issues arise with the estimation of aggregate separation rates from micro data. Moreover, much of the micro data required to even generate such estimates is not publicly available. In contrast, our results are solely based on publicly available macroeconomic data and are comparable across countries.

Our results can be summarized as follows. First, cross-country differences in job-finding rates are far more pronounced than differences in separation rates. Whereas in Anglo-Saxon and most Scandinavian countries, on average more than 10% of unemployed persons find a job in any given month, in most other countries in our sample this rate is 8% or lower. The U.S. seems to have by far the highest monthly job-finding rate with 56.3%, while Italy's 2.58% is the lowest in our sample. We estimate the fraction of workers that leave their jobs to be between 2.0% and 0.7% for the countries in our sample. Second, in spite of our different estimation method, our estimates are both qualitatively and quantitatively in line with those presented in previous studies (Shimer (2005), Hall (2005a,b), Fujita and Ramey (2006), Ridder and van den Berg (2003), and Jolivet, Postel-Vinay, and Robin (2006)). Finally, we combine our estimates with evidence on unemployment

[2] See Card, Chetty, and Weber (2007) for a detailed survey.

and labor force participation rates and impute steady-state worker flows between unemployment, employment, and not-in-the-labor-force for 23 of the countries in our sample.

2 Estimation

We use GMM to estimate the steady-state job-finding and separation rates for a broad sample of countries. In this section we first describe the theoretical moment conditions that are the foundation of our estimation of job-finding rates and then derive how we can use data on the unemployment duration distribution to construct the sample equivalent of these moments. We briefly discuss how data on the distribution of job tenure lengths can be used in a similar fashion to obtain estimates of average separation rates. The details underlying the derivations in this section are in the Appendix.

2.1 job-finding rates

In the steady-state of a continuous-time model of unemployment with a job-finding rate f_m that potentially varies across the length of the unemployment spell, m, the number of persons unemployed for m months, U_m, satisfies the differential equation

$$\dot{U}_m = -f_m U_m \tag{1}$$

such that

$$U_m = e^{-\int_0^m f_s ds} U_0, \tag{2}$$

were U_0 is the constant number of workers that flows into unemployment.

The total number of unemployed persons in this economy is

$$U = \int_0^\infty U_m dm. \tag{3}$$

The fraction of these workers that finds a job is

$$f = \frac{1}{U} \int_0^\infty f_m U_m dm = \frac{U_0}{U}, \tag{4}$$

which is the average job-finding rate across unemployed workers and is the job-finding rate that is of most interest to macroeconomists.[3] This is the rate that we aim to estimate for a broad set of

[3] For expository purposes, we call f_m the rate at which persons that have been unemployed for m periods find a job. In actuality, not all unemployed persons that leave unemployment do so because they find a job. Hence, more formally, f_m can be interpreted as the duration dependent hazard rate for leaving unemployment.

countries.

In practice, we do not have data on the whole distribution of unemployed workers across unemployment spells; we only have unemployment duration data in bins. We index these bins by $b = 1, \ldots, B$ and denote the shortest spell in each bin by m_{b-1} and the longest by m_b. We denote the fraction of unemployed persons in each bin by u_b, such that

(5) $$u_b = \frac{\int_{m_{b-1}}^{m_b} U_m dm}{U} = f \int_{m_{b-1}}^{m_b} e^{-\int_0^m f_s ds} dm$$

Because we only have a limited number of bins, we cannot estimate the full shape of the job-finding rate function, f_m. Instead, what we assume is that the duration dependence of the job-finding rate takes the form of a Gompertz hazard function, which only depends on two parameters. That is

(6) $$f_m = \lambda \exp(-gm), \text{ where } \lambda, g > 0$$

Here, λ reflects the initial job-finding rate at the moment of entry into unemployment and g is the per month percentage decline in that initial job-finding rate.

If the data were generated from the steady-state of the search model then the fractions u_b and the average job-finding rate, f, would be constant over time. This is not the case in the data, however, because economic fluctuations cause both the flow into unemployment, as well as the job-finding rate, to fluctuate over time. Moreover, the data on the distribution u_b is based on survey evidence and thus contains measurement error. For each country, we use exactly identified GMM to estimate the average monthly job-finding rate, f, using the moment condition that the distribution of unemployed workers over unemployment spells on average equals that in the steady-state.

Let $u_{b,t}$ be the observed fraction of workers in bin b at time t. We assume that

(7) $$E\left[1 - \frac{u_b}{u_{b,t}}\right] = 0 \text{ for all } b = 1, \ldots, B-1 \text{ and } t = 1, \ldots, T.$$

Because $\sum_{b=1}^{B} u_b = 1$ the B moment conditions are highly correlated,[4] we use only the first $B-1$ moment conditions.

These moment conditions allow us to estimate the Gompertz parameters λ and g using GMM with an identity-weighting matrix. This boils down to finding the parameter estimates $\widehat{\lambda}$ and $\widehat{g}$

[4] Since the moment conditions are defined in percentage deviations, the adding-up constraint does not necessarily imply perfect correlation between all B moment conditions. In practice, however, they turn out to be very highly correlated.

that minimize the sum of squared residuals,

$$\sum_{b=1}^{B-1}\left(1-\frac{\overline{u}_b}{u_{b,t}}\right)^2, \text{ where } \overline{u}_b=\frac{1}{T}\sum_{t=1}^{T}u_{b,t}. \tag{8}$$

Given our estimates $\widehat{\lambda}$ and $\widehat{g}$, we then use (4) to estimate the average job-finding rate. We denote this estimate by $\widehat{f}$. The reason that we use the identity-weighting matrix, as opposed to the efficient weighting matrix, is that we have at most about 20 years of observations in the different data sets that we use ($T < 20$ in our applications). Using the efficient-weighting matrix would therefore lead to a small sample bias in the results.

We calculate the standard error of the estimated average job-finding rate, denoted by $\widehat{\sigma}_{\widehat{f}}$, using a HAC estimated with a Bartlett kernel with bandwidth of 4. This allows for heteroskedasticity and serial correlation in the moment conditions due to the persistence in the effect of aggregate economic fluctuations on the duration distribution of unemployment spells. Because of the short time series that we use, estimated (asymptotic) standard errors that we report should be interpreted with caution. They most likely substantially underreport the actual (small sample) standard error of our analysis.

2.2 Separation rates

Let s_τ denote the rate at which workers that have been with the same job for τ periods leave that job. Then, in the steady-state of a continuous-time search model of unemployment, the number of persons that are employed with a job tenure of length τ, which we denote by E_τ, satisfies

$$\dot{E}_\tau = -s_\tau E_\tau. \tag{9}$$

The fraction of the employed that leave their jobs, which is the overall separation rate that we aim to estimate, is given by

$$s=\frac{1}{E}\int_0^\infty s_\tau E_\tau dm=\frac{E_0}{E}, \text{ where } E \text{ is the total number of employed people.} \tag{10}$$

Given binned data on the job tenure distribution of employees, we can estimate the aggregate separation rate in a way similar to the job-finding rate, above.

As with the duration dependent separation rates, we assume that the tenure-length-dependent separation rate, s_τ, has the shape of a Weibull hazard function and estimate both the initial separation rate and the rate at which the separation rate declines.

3 Data

Our main data source is the OECD (2006a,b) *Employment and Labour Market Statistics*. The countries we cover are Australia, Austria, Belgium, Canada, the Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Poland, Portugal, the Slovak Republic, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

Unemployment duration distributions

We take the unemployment duration distributions that we use to estimate the job-finding rates from the "incidence of unemployment duration" statistics from OECD (2006a). These data are reported in five duration bins: $<$ 1 month, $>$1 month and $<$3 months, $>$3 months and $<$6 months, $>$6 months and $<$1 year, 1 year and over. Moreover, for all countries, the data cover unemployed persons of age 15 and older. The years covered vary by country; the earliest year for any country is 1968, and the latest is 2004. Table 1 shows the sample period for each country, as well as the average duration distribution of the unemployed over the five duration bins over the sample period.

The cross-country variation in the average unemployment duration distribution is quite astounding. In the U.S. more than 70% of unemployed persons have been unemployed for less than one quarter, and only 7% have been unemployed for more than a year; however, in most continental European countries the distribution is almost the reverse, with 7% or less being in a spell of one quarter or shorter and more than one third of the unemployed having been so for more than a year.

Our identifying assumption is that, over the sample period, the unemployment duration distribution fluctuates around the average distribution reported and that there is no shift in this average distribution. This turns out not to be a bad approximation for most countries in our sample. However, there are notable exceptions; Denmark and the U.K., for example. In these countries labor market reforms have taken place during our sample period that have caused a persistent shift in the unemployment duration distribution. Kongshøj Madsen (1999) and Pissarides (2003) describe the labor market reforms that have led to this shift in the incidence of unemploymend duration in Denmark and Britain, respectively.

Job tenure distributions

We take job tenure distribution data from the "employment by job tenure intervals" statistics, from OECD (2006b). The OECD only publishes these statistics for EU member states. These data cover workers of age 15 and older. We have acquired data for non-EU countries from three country-specific datasources: Australia (Australian Bureau of Statistics, 2006), Canada (Statistics Canada, 2006), and the U.S. (Bureau of Labor Statistics, 2006). We consider the data in seven tenure length bins: <1 month, >1 month and <6 months, >6 months and <1 year, >1 year and <3 years, >3 years and < 5 years, >5 years and < 10 years, 10 years and over.

Table 2 shows the sample periods, as well as the average job tenure distribution, for each of the countries in our sample.[5] The job tenure data span a shorter time period than those on unemployment duration; they cover 1992 through 2006. The differences in job tenure distributions across countries are much less profound than those in unemployment duration. When we consider the percentage of workers that have held their job for more than a decade, however, we find that this is generally lower than 30% in the Anglo-Saxon countries, while it is 35% or higher in continental Western Europe. The extremes are, on the low end, Australia (24%) and the U.S. (26%) and, on the high end, Greece (52%) and Italy (49%).

Because job tenure distributions are not as much affected by business cycle fluctuations as unemployment duration distributions, the time variation in the former is much smaller than in the latter.

4 Results

We present our results in three parts. First, we focus on our estimates, considering their magnitude, and comparing them across countries and over time. Second, we compare our estimates for a subset of our sample of countries with those presented in other studies. Finally, we combine our estimates with average unemployment and labor force participation rates to obtain imputed job flows across different labor market states for the working age population.

4.1 Estimated job-finding and separation rates

Tables 1 and 2 contain the estimated job-finding and separation rates respectively. In both tables, $\widehat{\lambda}$ is the estimated initial hazard rate, $\widehat{g}$ is the estimated per month percentage decline in the hazard

[5] Because of the different data sources, not all data are provided in the same bins. For this reason, we report the percentage of workers with a job tenure in some combined bins for the three countries for which we use country-specific datasources.

rate, and $\widehat{\sigma}$ is the estimated standard error of the average hazard rate. These estimated average hazard rates are the monthly job-finding rate, $\widehat{f}$, in Table 1 and the monthly separation rate, $\widehat{s}$, in Table 2.

job-finding rates

The job-finding rates that we obtain vary substantially across countries. With 56.3%, the U.S. job-finding rate is by far the highest in the sample. In fact, it is more than 20 times higher than the estimated job-finding rate for Italy and is more than 8 times higher than that in continental Western European economies, like Belgium, France, Germany, the Netherlands, Spain, and Portugal.

Interestingly, the countries with relatively high job-finding rates have a substantial duration dependence of those rates over the first year of unemployment. For example, the job-finding rates at the moment of entry into unemployment in the U.S. is about 75.5%, as opposed to the cross-duration average of 56.3%. Similar differentials between the initial and average job-finding rates can be found for all countries for which we estimate average job-finding rates of 10% or higher.

For countries with low estimated average job-finding rates, like the continental Western European countries mentioned above, we find that the job-finding rate has almost no duration dependence. The unemployment duration distribution for these countries can be closely approximated by an exponential distribution (i.e. $g = 0$). Moreover, if the distribution is exponential, then the fraction of workers unemployed for less than a month can be used as a proxy for the monthly job-finding rate. For many countries with low estimated job-finding rates, we find that this approximation works reasonably well. In particular, the fractions are (with our estimates in parenthesis): France 6% (6.6%), Germany 7% (6.9%), Greece 6% (5.2%), the Netherlands 5% (4.7%), and Spain 4% (3.9%).

Our estimates can be interpreted as steady-state job-finding rates. Our identifying assumption is that this steady-state, or rather the average unemployment duration distribution on which our estimate is based, does not change over time. Because the estimation method requires several years of data, it is not always possible to assess whether such changes have occurred.

However, for Canada, France, Sweden, and the U.S., the data span a long enough period to allow for such an assessment. For these countries we calculated time series of estimated job-finding rates, based on 15 year rolling samples. The resulting time series are plotted in Figure 1. As can be seen from this figure, the estimated job-finding rates for France and Canada are essentially constant. The estimated job-finding rate in Sweden declines from 34% (1976-1990) to 20% (1990-2004). Similarly, the one in the U.S. declines from 63% (1968-1982) to 50% (1990-2004). In Sweden this decline in the job-finding rate was accompanied by a substantial increase in the unemployment rate. In the U.S., however, the opposite is the case.

The observed joint decline in the job-finding rate and in the unemployment rate in the U.S. is most likely due to structurally unemployed people making up a higher fraction of the pool of unemployed at the end of the sample than at the beginning. Since structurally unemployed workers face a job-finding rate of almost zero, an increase in their share in the pool of the unemployed will lead to a decrease in the average job-finding rate.

To see how well the identifying assumption of an average Weibull hazard rate fits the data, we present the actual and fitted unemployment duration distributions for France, Germany, the U.K., and the U.S. in Figure 2. The boxes reflect the distribution in the data; the top of the box is the maximum fraction of unemployed workers observed in each bin, the bottom is the minimum, and the line in the box the average. The dots represent the unemployment duration distributions implied by the estimated Weibull hazard functions. The fitted distributions characterize the shape of the unemployment duration distribution remarkably well. France, Germany, and the U.K. have high rates of long-term unemployment, while the U.S. labor market is characterized by much shorter unemployment spells. In particular, for France, Germany, and the U.K., around 40% of the unemployed experience more than a year of joblessness while in the United States more than 40 percent of unemployed experience less than one month of joblessness. This stark difference is captured very well by our estimation procedure, as can be seen in the comparison of the actual and fitted distributions.[6].

Separation rates

Although we find a very substantial variation in job-finding rates across countries, we do not find such a variation in separation rates. The separation rates that we obtain for the 23 countries in our sample are between 0.70% (Hungary) and 2.03% (Spain). Hence, where job-finding rates in the OECD seem to differ by a factor of 20 or more, separation rates seem to only differ by a factor of 3.

Moreover, the countries that are known for having more flexible labor markets and lower un-

[6] Additional results with different forms of the hazard rate function, which are not reported here, were very similar to the ones presented here.

employment rates (e.g. Australia, Canada, Denmark, the U.K., and the U.S.) do not stand out in terms of higher estimated separation rates. This is most likely because labor market flexibility disproportionately affects short tenure jobs, and our estimates are in large part determined by data on persons that hold a job for three years or longer. It suggests that the Weibull hazard function does not fully capture the tenure-length dependence of separation rates.[7]

To what extent the Weibull hazard function generates the average job tenure distribution in the data can be seen in Figure 3. This figure, structured similarly to Figure 2, shows that the Weibull hazard function captures the main shape of the job tenure distributions for France, Germany, and the U.K. The fit lies outside of the range of observations more than for the unemployment duration distributions, partly because the job tenure distributions tend to fluctuate less over time. For the U.S., however, the Weibull hazard function does not seem to be able to generate the job tenure distribution in the data. It underfits the fraction of workers with short job tenures, resulting in an underestimate of the U.S. separation rate.[8]

4.2 Comparison with other studies

We are not the first to estimate job-finding and separation rates for use in macroeconomic search models of unemployment. In order to consider how reasonable our estimates are, we compare them with estimates previously reported for a subset of countries from our sample.

Estimates of the (time series of the) U.S. job-finding rate have been published in several recent studies: Shimer (2005), Hall (2005a,b), and Fujita and Ramey (2006). Shimer (2005) computes the job-finding rate for the U.S. from 1948 to 2004 by using data on the number of unemployed workers and the number of short-term unemployed workers from the Current Population Survey. In his benchmark calculations he assumes, consistent with our interpretation of our estimate as a 'job-finding rate', that workers neither enter nor exit the labor force, but simply transit between employment and unemployment.[9] We take this limitation of Shimer's and our estimates into

[7] Unreported results obtained with different functional forms for the hazard function were qualitatively very similar to the ones reported here, however.

[8] Alternative estimates, using different functional forms for the hazard function, as well as different bins and moment conditions, yielded a similarly poor fit.

[9] Shimer (2005) also computes the job finding probability by taking into account the flows between three labor market states: unemployment, employment, and inactivity. He finds that both measures of job finding probability are highly correlated and that his measure, that ignores inactivity, which is comparable to our estimate, is higher,

account in the next subsection. Fujita and Ramey (2006) also use the CPS to compute the job-finding rate for the U.S. Our estimate is best interpreted as the sum of their unemployment to employment ($U \rightarrow E$) and unemployment to not in labor force ($U \rightarrow N$) flows. Hall (2005a) also uses CPS data to estimate job-finding rates. His estimates are somewhat lower than Shimer's and Fujita and Ramey's since he includes discouraged workers and marginally attached workers as unemployed.

Figure 4 plots our estimated time series for the U.S. job-finding rate, which is the same as the one in Figure 1, alongside comparable estimates from Shimer (2005), Hall (2005a,b), and Fujita and Ramey (2006). As can be seen from the figure, our estimated average U.S. job-finding rate of 56% is consistent with the estimates of Shimer (2005) and Fujita and Ramey (2006) and, as expected, slightly higher than that of Hall (2005a).

Cross-country comparable estimates of job-finding rates are few and far between. Ridder and van den Berg (2003) is probably most similar, in spirit, to our study, even though their main goal is to estimate an index of search frictions for five countries. They define the index of search frictions as the number of job offers that a worker receives during a spell of employment. They provide estimates of this index with micro, as well as aggregate, data. They argue that estimates obtained from aggregate data can be useful when micro data are not available or suffer from small sample problems.

Ridder and van den Berg (2003) report the average monthly flow out of unemployment as a percentage of unemployment from 1983-1995 for five countries.[10] Our estimates and theirs are very similar, even though we consider different time periods and use different estimation methods. In particular, their estimates and our estimates (in parentheses) are: France 3.6% (6.7%), Germany 7.6 % (7.0%), the Netherlands 6.6% (4.7%), the U.K. 7.7% (11.3%), and the U.S. 39.4% (56.3%). We consider these estimates to be quite similar given the different time periods, data sources, and estimation techniques. The correlation coefficient between these estimates, which is mainly driven by the higher U.S. job-finding rate in both studies, is 0.99.[11]

Shimer (2005) and Hall (2005) also estimate the U.S. separation rate. Shimer (2005) computes the entry rate to unemployment and finds that this rate averages around 3.4% from 1951 to 2003. Hall

since all flows out of unemployment are attributed to finding a job.

[10] See Table 1 in Ridder and van den Berg (2003) for more detail.

[11] For the four countries without the U.S., it is 0.37.

(2005b) computes the separation rate by using data from the Job Openings and Labor Turnover Survey (JOLTS) and finds that more than three percent of workers depart from employment each month. In another calculation, Hall (2005a) computes the separation rate by directly using the flows reported in the CPS and finds a separation rate of about seven percent. These estimates, combined with the fit plotted in Figure 3, suggest that our results probably reflect an underestimation of the U.S. separation rate.

As for cross-country comparisons of separation rates, Jolivet, Postel-Vinay, and Robin (2006) estimate job-spell hazard rates for eleven countries. They use European Community Household Panel Survey (ECHP) for the European countries and Panel Study of Income Dynamics (PSID) for the U.S. Their estimates are highly correlated with ours.

In particular, they find that Denmark, Ireland, Spain, and the U.K. have higher job-spell hazard rates than the other countries in their sample.[12] These countries also exhibit a clear negative duration dependence. For these four countries, our separation rate estimates are on the high side: Denmark 1.87 %, Ireland 1.39%, Spain 2.03% and the U.K. 1.53 % (All these estimates are on the higher side of our separation rate estimates). We also observe negative duration dependence for these countries. According to Jolivet, Postel-Vinay, and Robin's estimates, Belgium, France, Germany, Italy, the Netherlands, and Portugal have low job hazard rates. Our separation rate estimates for these countries are 0.92%, 1.14%, 1.06%, 0.69%, 0.99%, and 0.96%. In both their estimates and ours the U.S. lies in the middle.

In addition to Jolivet, Postel-Vinay, and Robin (2006), OECD (1997) contains estimates of separation rates for a set of OECD countries for two time periods, one in 1980s and one in 1990s, both of which represent periods of economic downturns for the countries in the sample. Nickell and Nunziata (2001) also report estimates of separation rates. Their estimates are the ratio of short term, less than a month, unemployment to the total number of employed persons. We compare our estimates with their average separation rates for 1992-1999. Figure 5 compares our estimated separation rates with those of the OECD and of Nickell and Nunziata (2001). The latter is denoted as the ratio of short term unemployment to employment in the figure. The OECD estimates are higher, maybe because they are measured during downturns, but our measure of separation rates is highly correlated with the other two. The correlation in both cases is 0.77.

Hence, the magnitude of our estimated separation rates are comparable to Jolivet, Postel-

[12] See Figure 1 in Jolivet, Postel-Vinay, and Robin (2006) for more detail.

Vinay, and Robin (2006), but are lower than the ones estimated for the U.S. using CPS data and the rates the OECD reports for economic downturns. This is probably due to the nature of the job tenure data on which our estimates are based. Job tenure data are often considered crude and relatively less reliable. Among the many problems associated with job tenure data in the literature are: inconsistency of reporting across calendar years, recall and rounding errors, spikes in the tenure distributions at years which are multiples of five (see Brown and Light, 1992). The relative separation rates that we estimate do seem to be very similar with those in Jolivet, Postel-Vinay, and Robin (2006) as well as those in OECD (1997).

4.3 Imputation of job flows

The estimated job-finding and separation rates above reflect the flow out of unemployment and the flow out of a job. These rates themselves are of use for the analysis of search models of the labor market. However, they do not capture the entire set of worker flows, nor the distribution of persons over labor market states, both of which are part of the steady-state equilibrium of such models. A full description of the labor market dynamics in such search models involves transition probabilities between the three labor market states: employed (E), unemployed (U), and not-in-the-labor-force (N).

Let f_{ij} be the fraction of workers that is in state i and flows to state j in a month. Then our estimate $\hat{f}$ can be decomposed into flows from U to E and U to N:

$$\hat{f} = f_{UE} + f_{UN}. \tag{11}$$

Similarly, our estimate $\hat{s}$ is the sum of flows from E to U, to N, and to another job, E':

$$\hat{s} = f_{EU} + f_{EN} + f_{EE'} \tag{12}$$

where $f_{EE'}$ denotes job-to-job transitions.[13] Thus, our estimated job-finding and separation rates are only an aggregate estimate across five of the ten possible labor market flows. All ten possible flows are depicted in Figure 6.

The aim of this section is to construct a set of imputed worker flows for all possible labor market transitions for the countries in our sample. We start of by presenting evidence on U.S. worker

[13] Our estimates, $\widehat{f}$ and $\widehat{s}$, are continuous time estimates. In the exposition and subsequent calculations we ignore time aggregation issues.

flows based on Fujita and Ramey (2006). Subsequently, we combine our estimates with cross-country evidence on unemployment and labor force participation rates, as well as three identifying assumptions on relative job flows, based on U.S. evidence, to impute the ten worker flows for 23 of the countries in our sample.

We use the labor market flows data for the U.S. compiled by Fujita and Ramey (2006) to examine the relative magnitude of labor market flows. The flow data that we use from Fujita and Ramey (2006) cover the flows f_{EU}, f_{EN}, f_{UE}, and f_{UN} from 1976 to 2005. We use these four flows and combine them with data on the number of employed (E_t) and unemployed (U_t) persons, as well as the number of people not-in-the-labor-force (N_t). These three stocks evolve according to the following difference equations:

$$\begin{aligned} E_t &= E_{t-1} + f_{UE,t}U_{t-1} + f_{NE,t}N_{t-1} - f_{EU,t}E_{t-1} - f_{EN,t}E_{t-1} \\ U_t &= U_{t-1} + f_{EU,t}E_{t-1} + f_{NU,t}N_{t-1} - f_{UN,t}U_{t-1} - f_{UE,t}U_{t-1} \end{aligned} \tag{13}$$

In addition, the transition probabilities satisfy the following adding-up constraints

$$1 = (f_{EE} + f_{EE'}) + f_{EU} + f_{EN} = f_{UE} + f_{UU} + f_{UN} = f_{NE} + f_{NU} + f_{NN}, \tag{14}$$

which hold for any t. The above five equations allow us to impute the flows $(f_{EE} + f_{EE'})$, f_{UU}, f_{NE}, f_{NU}, and f_{NN}.[14] The resulting time series of annual average monthly U.S. worker flows are listed in Table 3.

This imputation does not allow us to distinguish between job stayers and job-to-job transitioners; however, according to evidence from Fallick and Fleischman (2004) and Nagypál (2005), job-to-job transitions make up 40% of all flows out of a job. In terms of our worker flows, this implies that

$$f_{EE'}/(f_{EU} + f_{EN} + f_{EE'}) = 0.4. \tag{15}$$

We use this equation to impute job-to-job flows based on the results in Table 3.

Unfortunately, labor market flow data, like that in Fujita and Ramey (2006), are not available for most countries in our sample. Therefore, we resort to using our estimated job-finding and separation

[14] Fujita and Ramey (2006) provide similar estimates of $(f_{EE} + f_{EE'})$, f_{UU}, f_{NE}, f_{NU}, and f_{NN}. The similarity between our estimates and those of Fujita and Ramey (2006) result from their use of the margin error adjustment procedure of their flow data; this procedure implies that (13) holds by approximation for their flows.

rates, as well as cross-country evidence on unemployment and labor force participation rates taken from OECD (2006c), to impute steady-state worker flows for the countries in our sample.

Let U denote the steady-state fraction of persons in the working age population that are unemployed, E the fraction that are employed, and N the fraction that are not-in-the-labor-force. Given this notation, the unemployment rate is given by $U/(U+E)$, and the labor force participation rate is given by $(U+E)$. The steady-state unemployment and labor force participation rates are determined by the steady-state version of (13), which reads

$$\begin{aligned} (f_{EU}+f_{EN})\,E &= f_{UE}U+f_{NE}N \\ (f_{UN}+f_{UE})\,U &= f_{EU}E+f_{NU}N \end{aligned} \tag{16}$$

In conjunction with the definition of the job-finding rate, (11), the definition of the separation rate, (12), and the three adding up constraints, (14), this gives us seven equations. In order to identify the worker flows, we make three additional assumptions about the relative importance of particular worker flows.

First, to break down the job-finding rate, we assume that workers flowing out of unemployment are equally likely to go to employment or not-in-the-labor-force. That is,

$$f_{UE}/f_{UN}=1 \tag{17}$$

This is consistent with the evidence presented in Table 3, which shows this ratio is around 1.4 for the U.S. Fallick and Fleischman (2004) report this same ratio as 1.2 for the U.S., while for New Zealand it is, on average, 1.03.[15]

Second, to decompose the separation rate, we use the estimates of Fallick and Fleischman (2004) and Fujita and Ramey (2006). Their estimates for the U.S. imply that 66% of flows out of employment go to not-in-the-labor-force. This translates to

$$f_{EN}/f_{EU}=2 \tag{18}$$

Finally, we assume that (15) applies to all countries in our sample.

Of course, these three assumptions come from the U.S. labor market flows and are likely to be different in other countries. Therefore, we consider our imputations as, at best, a first step in calibrating macroeconomic models of labor markets for cross-country comparisons.

[15] Taken from the Household Labour Force Survey for 1994-2006. See *Statistics New Zealand* (2006).

Table 4 reports the imputed worker flows. The last line contains the average U.S. flows implied by Fujita and Ramey (2005). Because of the relatively low estimated separation rates relative to job-finding rates, our imputed worker flows from N to E are sometimes negative. Where this is the case, we have set them equal to zero.

5 Conclusion

Although search models of the labor market have become common in macroeconomic analysis of unemployment, cross-country evidence on the search frictions that are at the heart of these models is very sparse. In this paper, we provide a set of cross-country comparable estimates of job-finding and separation rates for over 20 OECD countries.

While our estimates are solely based on publicly available macroeconomic data sources, they are still consistent with evidence for the U.S. and a small set of other countries based on micro data. Our results suggest that the cross-country variation in job-finding rates is much higher than that in separation rates. From this, one has to conclude that, in order for modern search theories of unemployment to explain cross-country differentials in labor market outcomes, the challenge is to find what underlies differences in job-finding rates, rather than those in separation rates.

If, in addition to the estimated job-finding and separation rates, we assume that the relative magnitudes of certain U.S. labor market flows apply to all countries in our sample, we can impute steady-state worker flows. We report these imputed flows for persons that are employed, unemployed, and not-in-the-labor-force. We are aware that these imputations are based on relatively crude assumptions and anticipate that they will be refined in the future by comparable cross-country evidence based on micro data.

Until that data becomes available, the estimates in this paper provide some useful guidance for the cross-country calibration of macroeconomic models of unemployment.

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A Mathematical details

Derivation of equation (4)

The average job-finding rate is

$$f = \int_0^\infty f_m \frac{U_m}{U} dm = \frac{1}{U}\int_0^\infty -\left(\frac{\partial \ln U_m}{\partial m}\right) U_m dm \tag{19}$$

$$= \frac{1}{U}\int_0^\infty -\left(\frac{\partial U_m}{\partial m}\right) dm = \frac{U_0}{U}\left[1-0\right] = \frac{U_0}{U} \tag{20}$$

$$= \frac{1}{\int_0^\infty e^{-\int_0^m f_s ds} dm} \tag{21}$$

Derivation of equation (5)

The fraction of workers in bin b is given by

$$u_b = \frac{1}{U}\int_{m_{b-1}}^{m_b} U_m dm = \frac{U_0}{U}\int_{m_{b-1}}^{m_b} e^{-\int_0^m f_s ds} dm = f\int_{m_{b-1}}^{m_b} e^{-\int_0^m f_s ds} dm \tag{22}$$

Estimation of the standard error

The reported standard error of the average job-finding rate is calculated using the Delta-method[16]. First of all, we estimate the long run covariance matrix of the moment conditions to determine the degree of uncertainty about each of the conditions. When we define

$$\mathbf{u}_t = \begin{bmatrix} u_{1,t} & \ldots & u_{B-1,t} \end{bmatrix}' \text{ and } \mathbf{u} = \begin{bmatrix} u_1 & \ldots & u_{B-1} \end{bmatrix}' \tag{23}$$

This long run covariance matrix is defined as

$$\mathbf{S} = \sum_{i=-\infty}^{\infty} \mathbf{\Gamma}_i, \text{ where } \mathbf{\Gamma}_i = E\left[\mathbf{u}_t \mathbf{u}_{t-i}'\right] \tag{24}$$

and we estimate it using

$$\widehat{\mathbf{S}} = \widehat{\mathbf{\Gamma}}_0 + \sum_{i=1}^{h-1} \frac{h-i}{h}\left(\widehat{\mathbf{\Gamma}}_i + \widehat{\mathbf{\Gamma}}_i'\right) \tag{25}$$

where the autocovariance matrices are estimated using

$$\widehat{\mathbf{\Gamma}}_j = \frac{1}{T}\sum_{t=i+1}^{T} \left(\mathbf{u}_t - \overline{\mathbf{u}}\right)\left(\mathbf{u}_{t-i} - \overline{\mathbf{u}}\right)', \text{ where } \overline{\mathbf{u}} = \frac{1}{T}\sum_{t=1}^{T} \mathbf{u}_t \tag{26}$$

and we have used the bandwidth $h = 4$. Let the vector with Gompertz parameters be given by

$$\boldsymbol{\theta} = \begin{bmatrix} \lambda & g \end{bmatrix}' \tag{27}$$

Then we estimate the standard error of the estimate average job-finding rate by

$$\widehat{\sigma}_{\widehat{f}} = \sqrt{\frac{1}{T}\frac{\partial f}{\partial \boldsymbol{\theta}'}\left[\frac{\partial \mathbf{u}}{\partial \boldsymbol{\theta}'}\right]^{-1} \widehat{\mathbf{S}} \left[\frac{\partial \mathbf{u}'}{\partial \boldsymbol{\theta}}\right]^{-1} \frac{\partial f}{\partial \boldsymbol{\theta}}} \tag{28}$$

[16] See, for example, Hayashi (2000) for a derivation of the asymptotic standard error of GMM estimators.

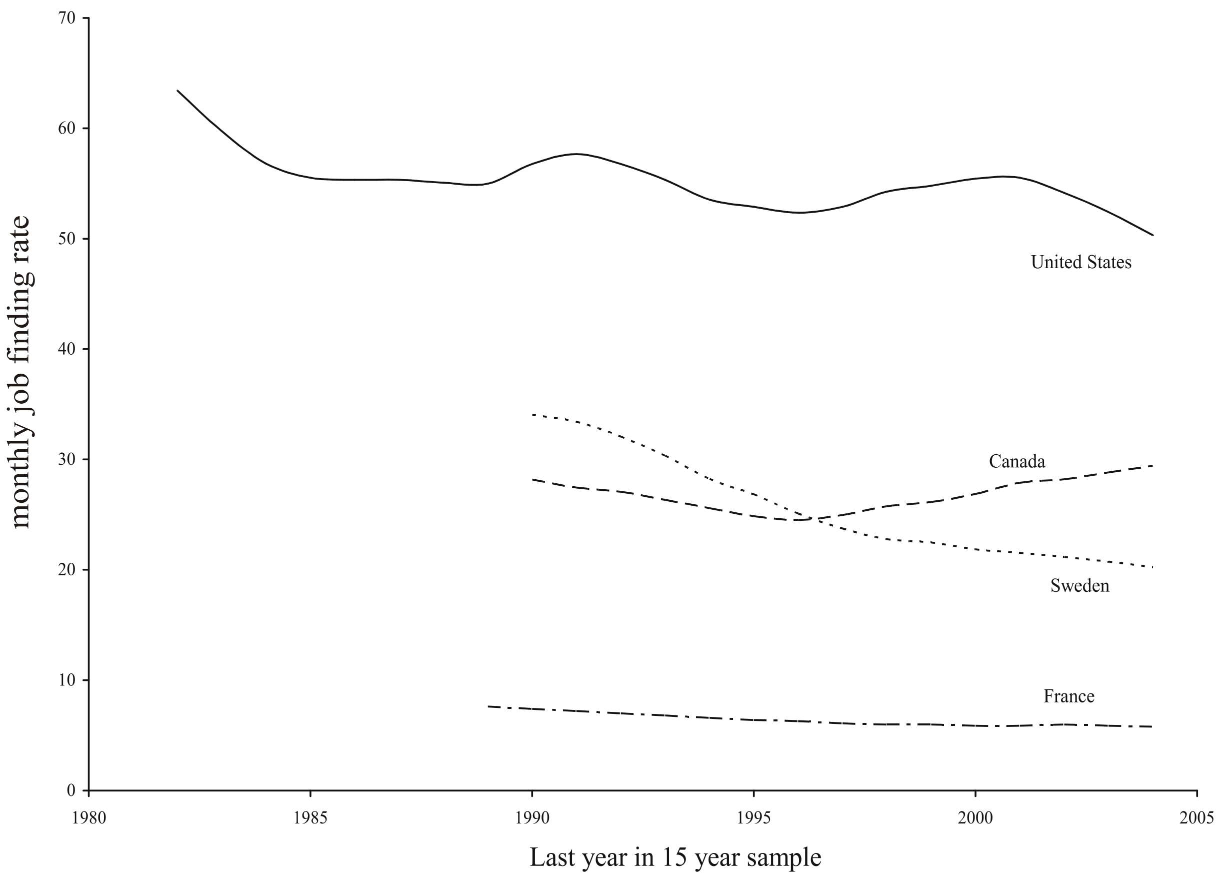


Figure 1: Job finding rates estimated based on a rolling 15 year sample

Table 1: Estimation results for job-finding rates

	Sample		Average unemployment duration distribution								
Country	**Start**	**End**	**0-1**	**1-3**	**3-6**	**6-12**	**>12**	$\widehat{\lambda}$	$\widehat{g}$	$\widehat{f}$	$\widehat{\sigma}_{\widehat{f}}$
Australia	1978	2004	18	22	14	20	26	26.90	4.54	17.05	1.15
Austria	1994	2003	12	30	18	15	25	24.70	4.18	15.61	0.74
Belgium	1983	2004	6	6	11	15	62	3.60	0.14	3.45	0.71
Canada	1976	2004	22	36	19	13	10	40.10	6.46	28.90	1.86
Czech Republic	1993	2004	9	15	17	21	38	9.80	1.30	8.06	1.60
Denmark	1983	2004	16	13	24	20	27	12.80	1.94	9.64	1.67
Finland	1995	2004	13	20	21	18	29	19.30	3.14	13.36	0.96
France	1975	2004	6	16	18	22	38	6.70	0.02	6.69	0.97
Germany	1983	2004	7	11	17	18	47	9.50	1.48	6.98	0.24
Greece	1983	2004	6	8	15	21	50	5.30	0.02	5.28	0.66
Hungary	1992	2004	7	12	15	22	45	6.50	0.10	6.41	0.69
Iceland	1991	2004	31	23	22	11	13	45.60	7.44	30.47	3.40
Ireland	1983	2004	5	9	13	19	54	4.00	0.02	3.98	1.87
Italy	1983	2004	4	5	12	17	63	2.60	0.02	2.58	1.72
Japan	1977	2004	18	27	17	19	19	26.80	4.34	19.07	1.06
Luxembourg	1983	2004	12	13	23	22	29	10.30	1.36	8.51	1.37
Netherlands	1983	2004	5	11	14	24	45	4.70	0.02	4.68	2.86
New Zealand	1986	2004	22	24	17	17	20	31.40	5.14	21.71	2.36
Norway	1983	2004	31	26	16	14	13	45.20	7.36	30.53	3.71
Poland	1992	2004	7	12	16	24	41	7.20	0.02	7.20	1.08
Portugal	1986	2004	7	14	19	16	44	7.10	1.46	3.88	1.36
Slovak Republic	1994	2004	6	10	13	18	54	7.70	1.22	5.65	0.32
Spain	1977	2004	4	15	15	19	48	4.00	0.02	3.98	1.54
Sweden	1976	2004	25	25	19	15	16	35.60	5.78	25.17	2.81
Switzerland	1991	2004	13	20	20	19	28	18.90	3.04	13.35	0.93
United Kingdom	1983	2004	12	17	18	17	37	17.60	2.98	11.27	1.53
United States	1968	2004	42	31	13	7	7	75.50	11.58	56.30	3.13

Note: The distribution, estimated coefficients and job-finding rates, and standard errors are all reported in percentages.

Table 2: Estimation results for separation rates

Country	Sample		Average job tenure length distribution										
Country	**Start**	**End**	**0-1**	**1-6**	**6-12**	**12-36**	**36-60**	**60-120**	**>120**	$\widehat{\lambda}$	$\widehat{g}$	$\widehat{s}$	$\widehat{\sigma}_{\widehat{s}}$
Australia	1992	2006	⊢ 13 ⊣		8	22	14	18	24	2.52	0.50	1.75	0.03
Belgium	1992	2004	2	4	5	13	11	18	46	1.68	0.49	0.92	0.04
Canada	1992	2006	⊢ 14 ⊣		8	⊢ 31 ⊣		17	30	2.76	0.59	1.78	0.02
Czech Republic	1997	2004	2	4	6	15	16	27	30	0.96	0.02	0.94	0.04
Denmark	1992	2004	4	8	9	17	13	17	32	3.39	0.81	1.87	0.06
Finland	1995	2004	6	6	7	12	10	18	41	3.39	1.01	1.38	0.05
France	1992	2004	3	5	6	13	11	18	44	2.37	0.70	1.14	0.03
Germany	1992	2004	3	4	7	15	12	19	40	1.77	0.46	1.06	0.03
Greece	1992	2004	2	3	3	11	11	18	52	1.20	0.37	0.70	0.03
Hungary	1997	2004	2	4	6	15	16	24	33	1.14	0.13	0.99	0.05
Iceland	1995	2002	4	7	11	16	12	18	31	3.45	0.84	1.85	0.07
Ireland	1992	2004	3	6	8	16	13	17	37	2.52	0.65	1.39	0.11
Italy	1992	2004	2	3	4	12	11	19	49	1.02	0.26	0.69	0.07
Luxembourg	1992	2004	2	3	5	15	13	21	42	1.11	0.23	0.82	0.03
Netherlands	1992	2004	3	4	6	18	13	19	36	1.50	0.36	0.99	0.10
Norway	1995	2004	4	5	7	17	13	18	36	2.34	0.59	1.34	0.05
Poland	1997	2004	3	4	6	12	12	20	43	1.77	0.50	0.99	0.07
Portugal	1992	2004	3	4	5	13	11	18	45	1.80	0.53	0.96	0.04
Spain	1992	2004	6	10	8	12	10	16	38	5.70	1.57	2.03	0.16
Sweden	1995	2004	4	5	5	14	11	18	43	1.56	0.46	0.87	0.12
Switzerland	1996	2004	3	5	8	15	14	20	35	1.86	0.44	1.19	0.03
United Kingdom	1992	2004	3	6	9	17	14	19	32	2.52	0.59	1.53	0.04
United States	2000	2006	⊢ 15 ⊣		9	13	17	19	26	1.71	0.43	1.06	0.02

Note: The distribution, estimated coefficients and separation rates, and standard errors are all reported in percentages.

Table 3: US job market flows

Year	E to E&E'	E to U	E to N	U to E	U to U	U to N	N to E	N to U	N to N
1976	94.6	1.8	3.6	22.3	51.7	26.1	5.7	3.3	91.0
1977	94.8	1.7	3.5	23.1	48.8	28.1	5.8	3.2	90.9
1978	94.9	1.6	3.5	24.5	45.1	30.4	6.1	3.1	90.8
1979	94.9	1.6	3.5	24.0	45.1	30.9	6.3	3.0	90.7
1980	94.8	1.8	3.4	22.0	50.0	28.1	5.7	3.4	90.9
1981	94.8	2.0	3.2	22.0	51.6	26.5	5.6	3.4	91.0
1982	94.5	2.4	3.1	20.1	57.4	22.5	5.3	3.7	91.0
1983	94.9	2.1	3.0	20.4	57.3	22.3	5.2	3.7	91.1
1984	95.0	1.8	3.2	23.4	51.1	25.5	5.6	3.5	90.9
1985	94.9	1.9	3.2	23.2	49.3	27.5	5.9	3.4	90.7
1986	95.0	1.9	3.1	22.2	49.8	28.0	6.1	3.2	90.6
1987	95.1	1.8	3.1	22.0	47.9	30.1	6.5	2.9	90.6
1988	95.2	1.7	3.1	21.8	47.1	31.0	6.6	2.6	90.8
1989	95.4	1.7	3.0	20.6	47.1	32.3	6.8	2.4	90.8
1990	95.4	1.7	2.9	19.1	50.2	30.6	6.6	2.4	91.0
1991	95.3	1.9	2.8	17.1	55.5	27.4	6.2	2.7	91.2
1992	95.4	1.9	2.8	15.9	58.4	25.7	6.3	2.8	90.9
1993	95.3	1.8	2.9	15.5	57.6	26.9	6.8	2.5	90.7
1994	95.3	1.6	3.0	18.1	52.1	29.8	6.9	2.7	90.4
1995	95.5	1.6	2.9	18.0	52.4	29.6	6.5	2.4	91.1
1996	95.7	1.5	2.8	18.1	52.9	29.0	6.6	2.3	91.1
1997	95.8	1.3	2.8	18.4	50.8	30.9	6.6	2.3	91.1
1998	95.9	1.3	2.8	20.3	47.4	32.3	6.3	2.3	91.4
1999	95.9	1.2	2.8	22.5	44.5	33.0	6.3	2.3	91.4
2000	95.8	1.2	2.9	22.9	43.6	33.5	6.6	2.2	91.2
2001	95.6	1.4	3.0	21.4	47.6	31.1	6.4	2.5	91.2
2002	95.5	1.5	3.0	20.0	54.2	25.8	6.2	2.5	91.3
2003	95.7	1.4	2.9	20.5	54.6	24.9	5.7	2.7	91.5
2004	95.7	1.3	2.9	21.0	52.6	26.4	5.7	2.6	91.6
2005	95.8	1.3	3.0	22.1	50.7	27.2	5.9	2.5	91.6

Note: Based on combining Fujita and Ramey (2006) with Household Survey Data from Current Employment Situation reports, using (13) and (14).

Table 4: Imputed worker flows across countries

	Sample		**Percentage**		**Transition probabilities**								
	start	**end**	$\frac{U}{U+E}$	$U+E$	**E to E&E'**	**E to U**	**E to N**	**U to E**	**U to U**	**U to N**	**N to E**	**N to U**	**N to N**
Australia	1992	2004	7.6	74.1	99.0	0.4	0.7	8.5	83.0	8.5	0.2	4.1	95.7
Belgium	1992	2004	8.1	63.4	99.4	0.2	0.4	1.7	96.6	1.7	0.5	0.5	99.1
Canada	1992	2004	8.8	77.0	98.9	0.4	0.7	14.5	71.1	14.5	0.0	9.9	90.1
Czech Republic	1997	2004	7.5	72.2	99.4	0.2	0.4	4.0	91.9	4.0	0.2	1.7	98.0
Denmark	1992	2004	6.2	80.9	98.9	0.4	0.7	4.8	90.4	4.8	2.8	1.7	95.5
Finland	1995	2004	11.0	73.9	99.2	0.3	0.6	6.7	86.6	6.7	0.0	5.0	95.0
France	1992	2004	10.8	68.0	99.3	0.2	0.5	3.3	93.3	3.3	0.1	1.8	98.1
Germany	1992	2004	8.5	71.8	99.4	0.2	0.4	3.5	93.0	3.5	0.4	1.6	98.0
Greece	1992	2004	9.8	63.5	99.6	0.1	0.3	2.6	94.7	2.6	0.0	1.2	98.8
Hungary	1997	2004	6.7	59.9	99.4	0.2	0.4	3.2	93.6	3.2	0.3	0.8	98.9
Iceland	1995	2004	3.1	89.2	98.9	0.4	0.7	15.2	69.5	15.2	4.4	5.9	89.6
Ireland	1992	2004	8.8	66.0	99.2	0.3	0.6	2.0	96.0	2.0	0.9	0.6	98.5
Italy	1992	2004	10.5	60.5	99.6	0.1	0.3	1.3	97.4	1.3	0.2	0.5	99.3
Luxembourg	1992	2004	2.8	63.3	99.5	0.2	0.3	4.3	91.5	4.3	0.5	0.4	99.1
Netherlands	1992	2004	4.9	72.5	99.4	0.2	0.4	2.3	95.3	2.3	1.0	0.3	98.6
Norway	1995	2004	4.0	81.2	99.2	0.3	0.5	15.3	69.5	15.3	0.1	5.3	94.6
Poland	1997	2004	15.9	66.9	99.4	0.2	0.4	3.6	92.8	3.6	0.0	3.2	96.8
Portugal	1992	2004	5.6	74.4	99.4	0.2	0.4	1.9	96.1	1.9	1.1	0.3	98.5
Spain	1992	2004	17.1	65.3	98.8	0.4	0.8	2.0	96.0	2.0	0.7	1.4	97.9
Sweden	1995	2004	7.3	80.1	99.5	0.2	0.3	12.6	74.8	12.6	0.0	8.6	91.4
Switzerland	1996	2004	3.4	82.9	99.3	0.2	0.5	6.7	86.7	6.7	2.0	1.6	96.4
United Kingdom	1992	2004	6.9	77.4	99.1	0.3	0.6	5.6	88.7	5.6	1.1	2.5	96.4
United States	1992	2004	5.5	79.1	99.4	0.2	0.4	28.2	43.7	28.2	0.0	14.0	86.0
Fujita and Ramey	1976	2005			*95.6*	*1.5*	*2.9*	*19.4*	*51.4*	*29.1*	*6.4*	*2.5*	*91.2*

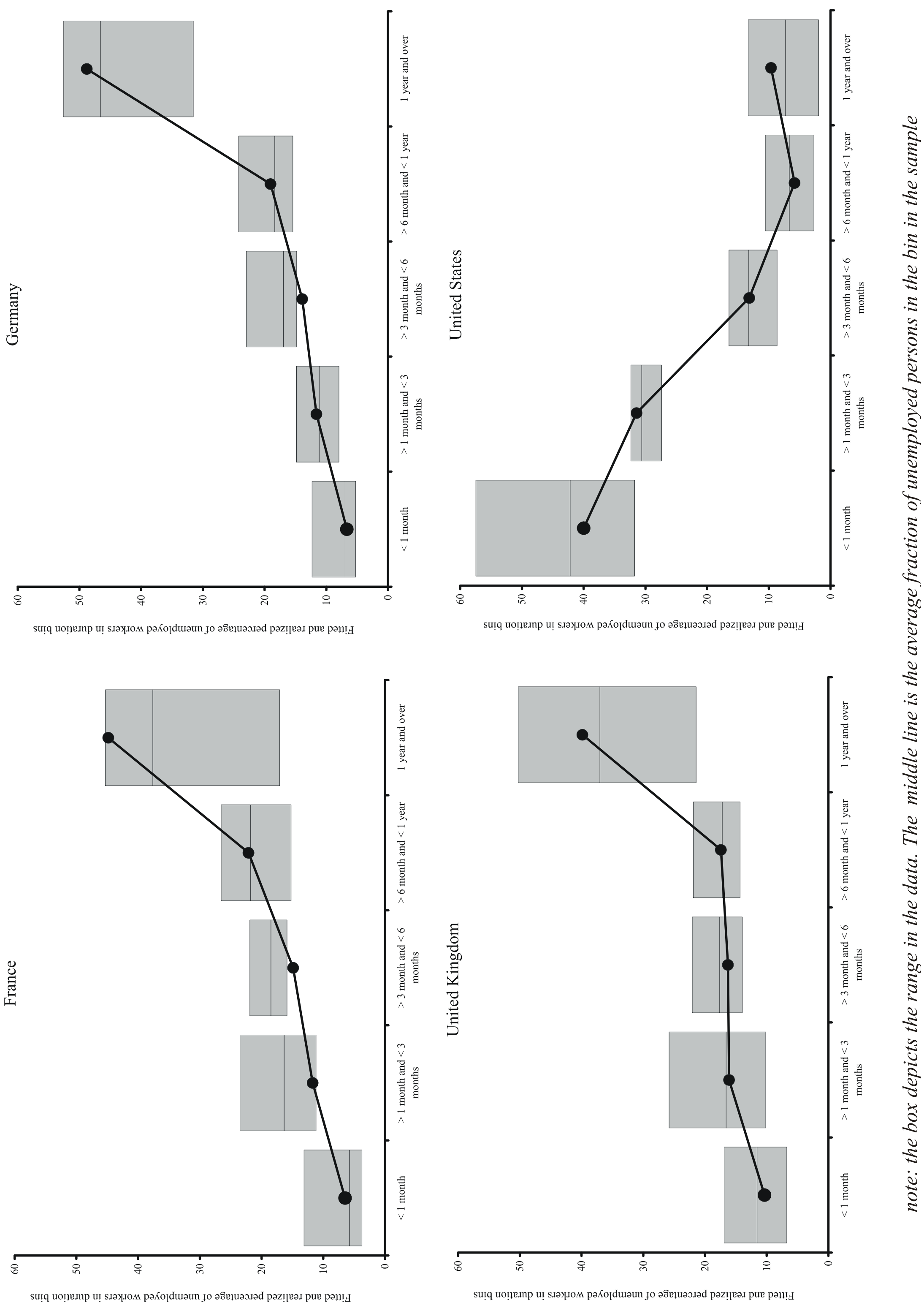


Figure 2: Actual and fitted unemployment duration distributions

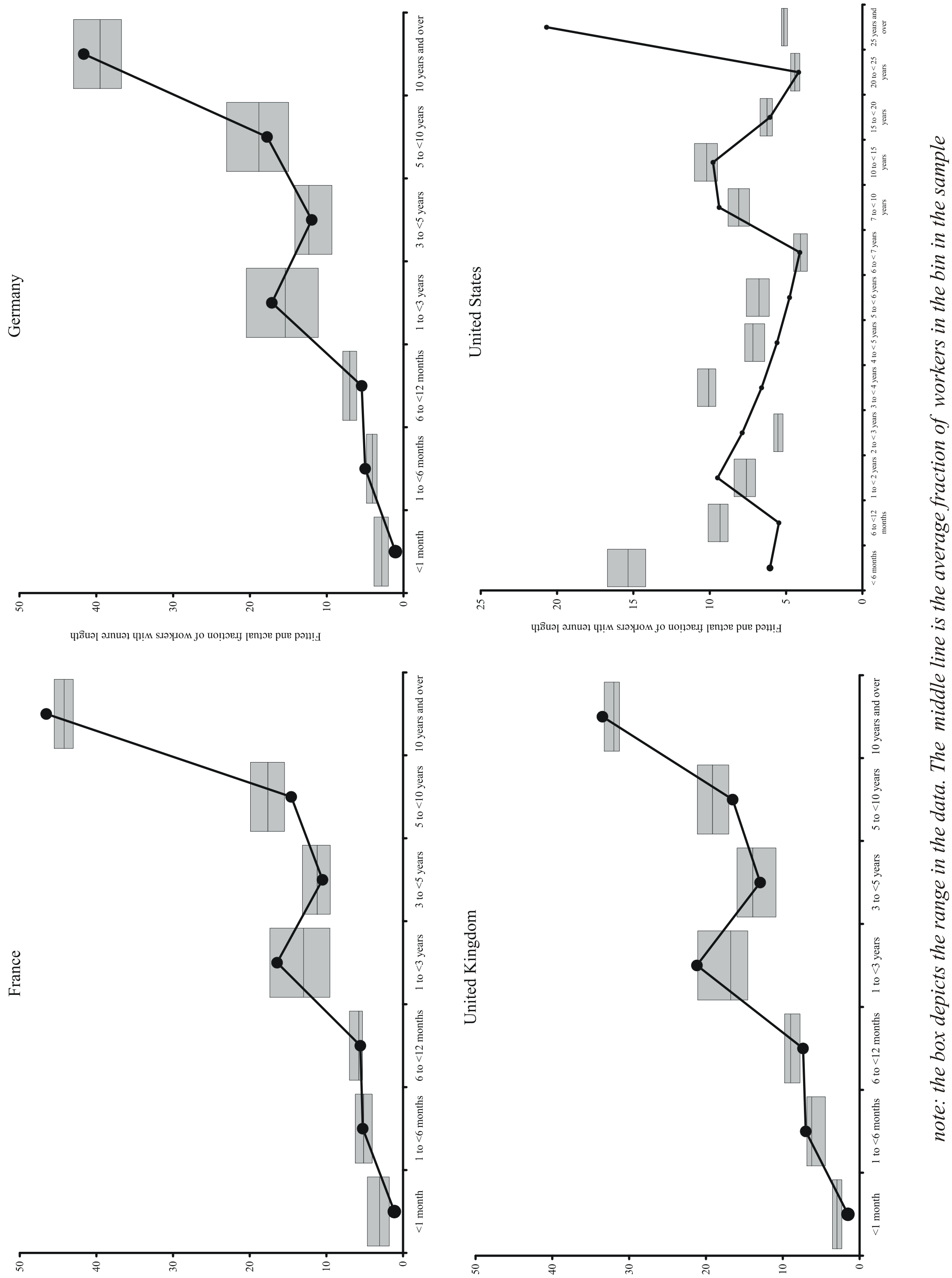
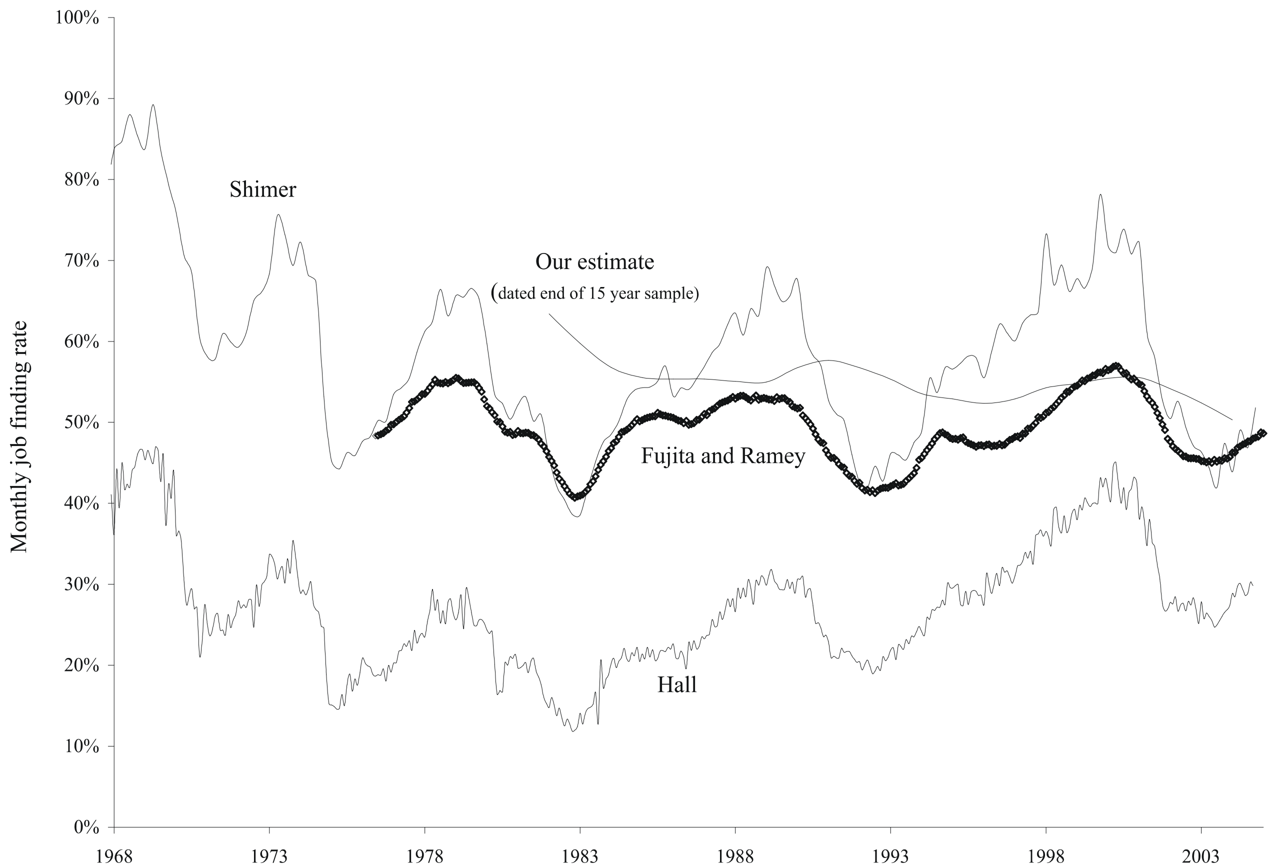


Figure 3: Actual and fitted job tenure distributions



Note:

Shimer: quarterly average of monthly job finding rate. This data was constructed by Robert Shimer. For additional details, please see Shimer (2005) and http://robert.shimer.googlepages.com/flows.

Fujita and Ramey: Annual average of monthly job finding rate from Fujita and Ramey (2006). We plot the sum of the transition rates from *U* to *E* and *U* to *N*, since this is consistent with our definition of the job finding rate

Hall: Ratio of new hires to the number of job-seekers, as in Hall (2005). Available at http://www.stanford.edu/~rehall/Selected_Publications.html

Figure 4: Comparison of estimates of U.S. job-finding rate.

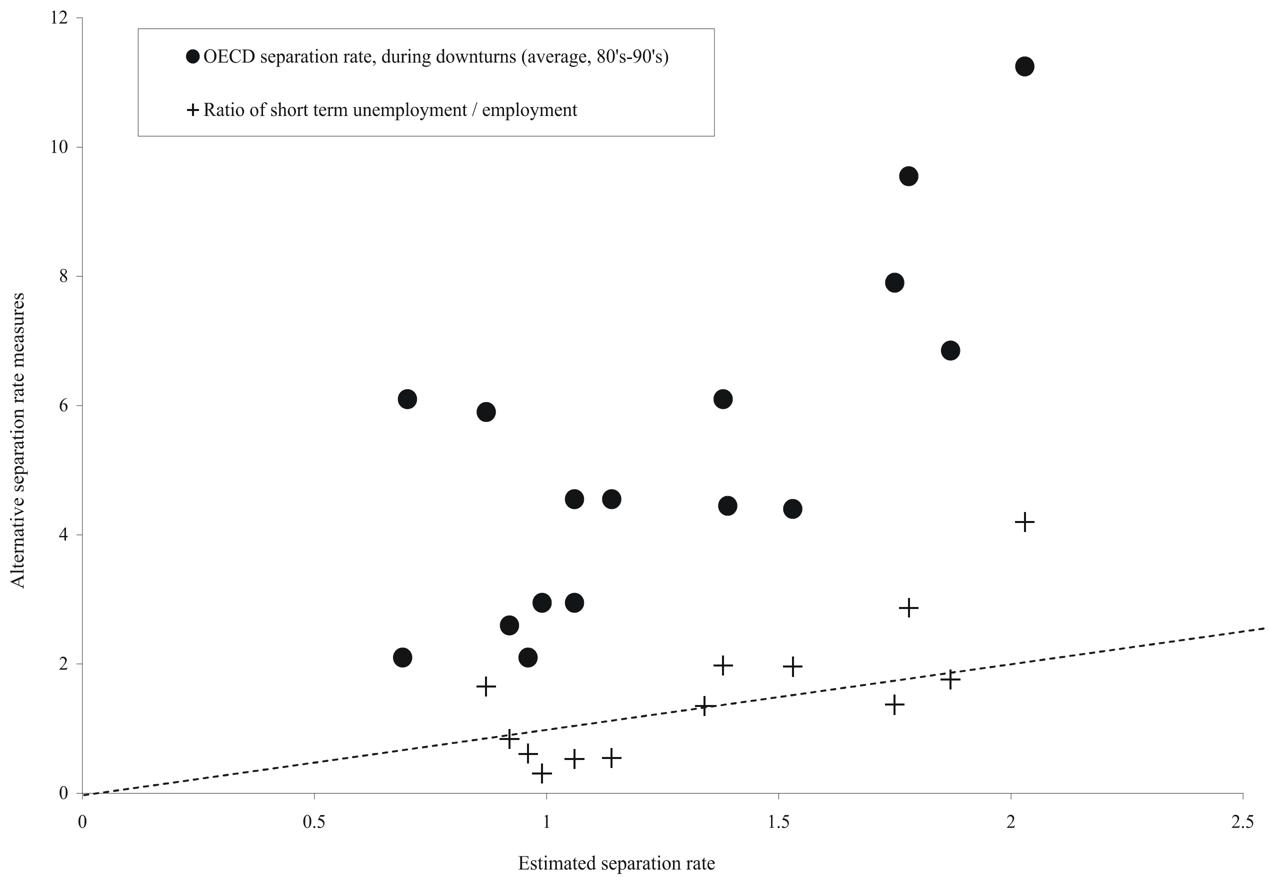


Figure 5: Comparison of our estimates with published OECD separation rates

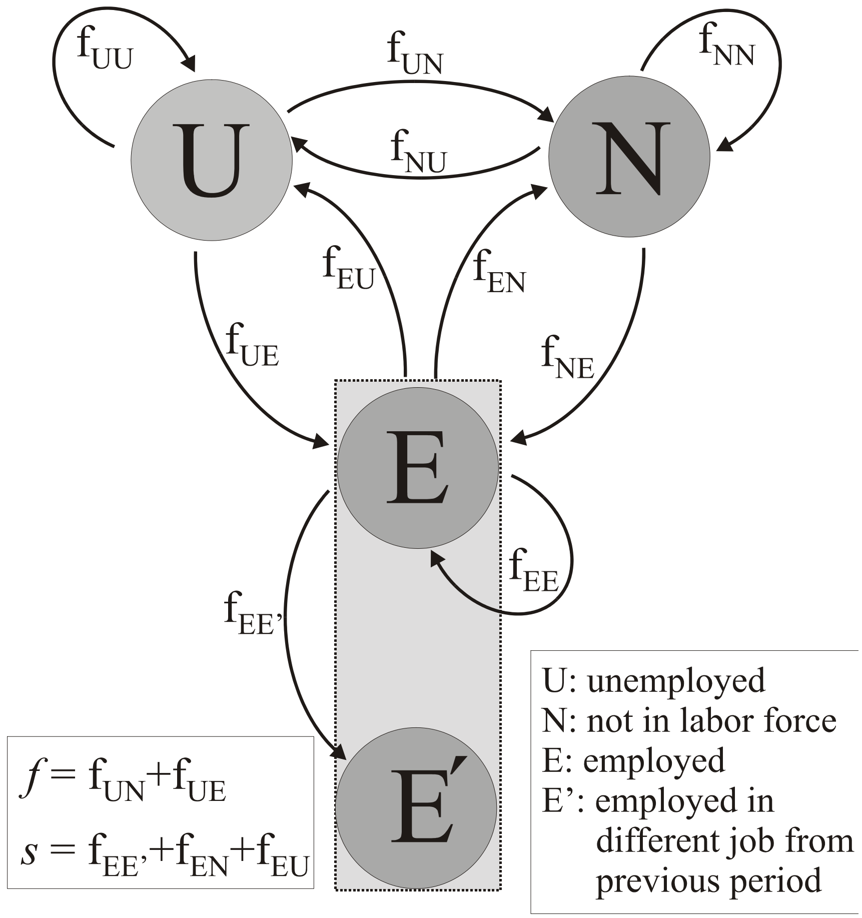


Figure 6: Worker flows diagram