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ARE INITIAL WAGE LOSSES OF INTERSECTORAL MOVERS COMPENSATED FOR BY THEIR SUBSEQUENT WAGE GAINS?

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This paper presents an equilibrium explanation of the inter- and intrasectoral mobility of workers. Analyses of our samples from the Panel Study of Income Dynamics and the National Longitudinal Survey of Youth show that, other things being equal, the initial wage decline is greater for intersectoral movers than for intrasectoral movers. Intersectoral movers, however, enjoy higher wage growth in subsequent years on postunemployment jobs than intrasectoral movers do, and hence are compensated for their initial wage decline. Our estimates suggest that, other things being constant, the additional short-term wage loss associated with sector shifts is overturned in no more than four years by the greater wage growth of intersectoral movers in subsequent years. The findings in the current study clearly show that the true economic costs of intersector mobility tend to be overstated in existing studies and are significantly lowered in the long-term perspective. Calibration of a simple lifetime utility model demonstrates that inter- and intrasectoral movements of workers are quantitatively consistent with an equilibrium framework, at least for a major group of workers who move with longer term perspectives. Evidence also shows that job seekers consider not only the initial wage rate but also the subsequent wages received from the postunemployment job when deciding whether to recommence employment or switch sectors.

Keywords: Intersectoral Mover, Intrasectoral Mover, Duration of Unemployment, Initial Wages, Wage Growth

1. INTRODUCTION

New technological change, severe recessions, and the resulting change in the composition of labor demand all induce workers to move from declining sectors to growing sectors. These sectoral shifts have attracted the attention of many

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researchers who want to explain unemployment fluctuations based on intersectoral labor mobility. For example, Lilien (1982) and subsequent papers from various authors have shown that the frequency of sectoral movers increases when the unemployment rate goes up.[1] Recently, Shin and Shin (2008) have reported that the lengthening duration of unemployment experienced by intersectorally mobile workers plays a major role in explaining cyclical unemployment.

Nevertheless, the question remains of why workers ever want to cross industrial lines upon job separation. A number of studies have confirmed that workers suffer from greater loss when they change sectors than when they stay in the same sector. For example, intersectoral movers (henceforth "movers") experience longer periods of unemployment than intrasectoral movers (henceforth "stayers")[2] Furthermore, other things being equal, the initial wage loss is greater for movers than for stayers [Addison and Portugal (1989); Carrington (1993); Houle and Audenrode (1995); Neal (1995)]. Why then do workers change sectors, when sector shifts are accompanied by various costs? This is particularly puzzling considering that, as emphasized by Davis and Haltiwanger (1990), most job reallocation is explained by intrasectoral rather than intersectoral reallocation.[3] Why do workers move across sectors, when a number of jobs are available in the same sector? Understanding the causes of intersector mobility may help us to understand unemployment fluctuations better.

Once a worker is separated from a job, voluntarily or involuntarily, he or she tries to find jobs in the industries that offer the best prospects in terms of search period and postunemployment wages. In the process, the worker might cross industrial lines for his/her own benefit. In fact, existing studies that report greater wage loss associated with sectoral shifts focus on the short-term wage consequences of sectoral shifts, while neglecting the long-term wage effects. We believe, however, that once the conventional job search framework encompasses a vision of long-term wage prospects, it can incorporate both long-term incentives that may favor sectoral shifts and short-term disincentives that make sectoral shifts less preferable.

The current study presents a basis for an equilibrium explanation of the sectoral mobility of workers. It focuses on the long-term incentives for intersectoral mobility and argues that movers may receive as many benefits as stayers do. Analyses of our sample from the Panel Study of Income Dynamics (PSID) show that, other things being equal, the initial wage loss of movers is approximately 14% greater than that of stayers. Movers, however, enjoy higher wage growth afterward than stayers do. The mover–stayer difference in the wage growth associated with an additional year of postunemployment tenure is estimated at 4.3%. These figures suggest that the additional short-term wage decline associated with sector change is overturned within four years by the higher wage growth of movers. The findings in the current study clearly show that the true economic costs of intersectoral mobility tend to be overstated in existing studies and are significantly lowered in the long-term perspective, insomuch as the current findings suggest that intersectoral mobility may be a rational choice for movers. Calibration of a simple lifetime utility model demonstrates that inter- and intrasectoral movements of workers can be quantitatively explained in an equilibrium framework, at least for a major group of workers who move with longer term perspectives.

Our empirical evidence also shows that the duration of unemployment and sectoral choice are jointly determined not just with the initial wage rate but also with the subsequent wages received from the postunemployment job. This suggests that although the duration of unemployment has long-term effects on wages possibly due to many factors such as the loss of general human capital, scar effects, and declining reservation wages, job seekers also consider the entire wage stream they expect to receive from a new job when deciding whether to recommence employment or whether to switch sectors.

The paper is organized as follows. Section 2 presents data and econometric methods, and Section 3 presents the empirical findings. On the basis of the estimates found in Section 3, Section 4 presents a quantitative explanation of inter- and intrasectoral mobility of workers in an equilibrium framework. Section 5 offers our conclusions.

2. DATA AND ECONOMETRIC METHOD

2.1. Data

We use the same sample analyzed by Starr-McCluer (1993) to obtain information on whether an individual's job change is made within the same sector or between different sectors and how long the intervening unemployment spell lasts. From 1981 through 1983, the PSID addressed detailed questions to household heads regarding various aspects of unemployment experienced during the previous year. Individuals were asked when a spell began and how many weeks it lasted. Upon reemployment by the survey week, the PSID also asked each individual the type of job change, that is, whether the new job was similar to or different from the old job. A person is defined as a "mover" when his or her new job is different from the old one. Otherwise, the person is classified as a "stayer."

This sample is suitable for the following reasons. First, as noted by Bound et al. (2001) and many other researchers, survey reports of industry affiliations of respondents are subject to great measurement error. In particular, the occurrence of changes in industry is exaggerated when estimates of such changes are obtained by comparing the reports on the industries obtained at two points in time. In contrast, the measurement error associated with the above-mentioned definition of sectoral mobility in the PSID sample is expected to be randomly made. Second, as emphasized by Starr-McCluer, the conventional measure of intersector change, that is, a change at the two-digit industry level, tends to exclude significant job changes within an industry but include trivial changes across industries. Third, ideally, we need to compare a stayer and a mover who are unemployed at the same point in time so that comparison of wage and unemployment experiences of the two workers is not contaminated by different aggregate economic environments pertaining to different years. At the same time, for practical purposes, we need to secure enough degrees of freedom for each type of worker, especially movers. Our sample meets these requirements, because, to be included in the sample, all sample spells must overlap the period between January 1980 and December 1982. This short period includes two recessions[4] and is expected to generate relatively large observations on movers. Fourth, as well recognized among many researchers, in general, the PSID sample represents the entire working age population better than other longitudinal data sets.

For each of the identified spells, wages on the postunemployment job are tracked by the year's survey week point until the person is separated from that job. This process requires information on the tenure of the postunemployment job. As noted by many researchers, tenure responses from PSID are often inconsistent with calendar time [Altonji and Shakotko (1987); Brown and Light (1992)]. To generate tenure variables that are internally consistent within a job stay, we borrow the complex algorithm developed by Altonji and Shakotko (1987), which has been applied in subsequent studies such as Solon et al. (1994).[5] We also observe the wages of preunemployment jobs recorded at the last survey week point.

Although our analysis is primarily based on the PSID, it is complemented by the National Longitudinal Survey of Youth (NLSY). This survey began in 1979 with a national sample of people aged between 14 and 22, who were then reinterviewed each year until 1994, and switched to biennial interviews thereafter. Despite the NLSY's restriction to this cohort and, therefore, its affording only a "second opinion," it is always desirable to double-check findings from the PSID with another longitudinal data set. In addition, the NLSY provides one other major advantage: because it makes a point of recording employer identifiers, it avoids the PSID's notorious difficulties with determining job tenure.

Although the particular PSID sample we consider here has clear information about intersectoral mobility, the NLSY sample lacks such information. Given the tendency toward overidentifying intersectoral mobility by comparing respondents' reports of the industries obtained at two points in time, we adopt the following two conservative approaches to classify inter- and intrasectoral mobility in analyzing the NLSY sample. In the first approach (hereafter Sector Classification 1), a job change is defined as intersectoral mobility if the last survey week report of the preunemployment job is different from the first survey week report of the postunemployment job in both industry and occupation at the *one*-digit level. A job change is defined as intrasectoral mobility when the last report accords with the first report in both industry and occupation at the *two*-digit level. Sector Classification 2 adopts an even more conservative standard: a job change is defined as intersectoral (intrasectoral) mobility if the last two consecutive industry reports of the preunemployment job coincide with each other at the two-digit level, the first two consecutive industry reports of the postunemployment job coincide with each other at the two-digit level, and the two industry codes are different (identical). Although Sector Classification 2 is expected to dramatically reduce measurement errors of the industry code within job duration, it tends to understate the volume of intersectoral mobility, as it eliminates all mobile workers whose either pre- or postunemployment job duration is less than two years. These workers are more likely to be movers, as job duration is shorter for movers than stayers (see Section 4).

2.2. Econometric method

To examine long-term wage consequences of job mobility, we express an individual's wage rate at a certain point of time as a function of his or her cumulative job mobility (all jobs as well as unemployment spells experienced ever since the start of the labor market career) along with other productive attributes,

$$\ln W_{i,m,t} = \sum_{h=1}^{m-1} \alpha_h \text{PastTenure}_{ih} + (\alpha_m + \beta_m)\text{CurrentTenure}_{i,m,t} + \sum_{h=1}^{n} \gamma_h \text{UnempDuration}_{ih} + \delta' Z_i + \eta' X_{it} + \varepsilon_{it}, \tag{1}$$

where $W_{i,m,t}$ is the ith individual's hourly rate of pay collected from his or her current job (denoted by job m) held during the survey week of year t, which is deflated by the Consumer Price Index (CPI) of that year; α is the transferable component of the return to job tenure; PastTenure_{ih} is completed job duration on past job h $\{h = 1, 2, 3, ...(m-1)\}$ held by individual i; β is the nontransferable component of the return to tenure; $\text{CurentTenure}_{i,m,t}$ is current tenure at time t; and $\text{UnempDuration}_{ih}$ is the completed duration of unemployment on spell h $(h = 1, 2, 3, ..., , n)$. Z_i represents a vector of all observable and unobservable individual-specific but time-invariant characteristics such as education, gender, race, ability, and motivation; X_{it} is a vector of time-varying individual characteristics; and ε_{it} is the error term.

Existence of unobservable components of Z_i and possible association of these with some included regressors often makes ordinary least squares (OLS) estimation of equation (1) inconsistent. An easy way of controlling for all unobservable as well as observable fixed effects is to consider a wage function at a different point of time and focus on wage changes between the two time points,

$$\ln W_{i,(m-1),o} = \sum_{h=1}^{m-2} \alpha_h \text{PastTenure}_{ih} + (\alpha_{m-1} + \beta_{m-1})\text{CurrentTenure}_{i,(m-1),0} + \sum_{h=1}^{n-1} \gamma_h \text{UnempDuration}_{ih} + \delta' Z_i + \mu' X_{io} + \varepsilon_{i0}, \tag{2}$$

where subscript 0 represents the last survey week point (end point) of the most recent job, the $(m-1)$th job, and time-varying characteristics are allowed to have different effects on the wage level between the current and the last job ($\mu \neq \eta$).

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Subtracting equation (2) from (1) yields

$$\ln {}^{W_{i,m,t}}\!/\!_{W_{i,(m-1),0}} = (\alpha_m + \beta_m)\text{CurrentTenure}_{i,m,t} - \beta_{m-1}\text{PastTenure}_{i,(m-1),0} + \gamma_n \text{UnempDuration}_{in} + \eta' X_{it} - \mu' X_{i0} + (\varepsilon_{it} - \varepsilon_{i0}). \quad \textbf{(3)}^{6}$$

In equation (3), all observable and unobservable individual fixed effects are "differenced out." To compare wage patterns of movers and stayers, we augment equation (3) by including a dummy variable, INTER_i (which equals one if person i changes sectors), and its interaction with the current tenure variable, $\text{INTER}_i \times \text{CurrentTenure}_{i,m,t}$:

$$\begin{aligned} \ln {}^{W_{i,m,t}}\!/\!_{W_{i,(m-1),0}} &= \lambda_0 \\ &+ \lambda_1 \text{INTER}_i + \lambda_2 \text{CurrentTenure}_{i,m,t} + \lambda_3 \text{INTER}_i \times \text{CurrentTenure}_{i,m,t} \\ &+ \gamma_n \text{UnempDuration}_{in} - \beta_{m-1}\text{PastTenure}_{i,(m-1),0} \\ &+ \eta' X_{it} - \mu' X_{i0} + (\varepsilon_{it} - \varepsilon_{i0}). \end{aligned} \quad \textbf{(4)}$$

Equation (4) is similar to the conventional specification adopted by existing studies such as Addison and Portugal (1989) that investigate the effect of the duration of unemployment on the wages of postunemployment jobs. However, there is one important difference. By extending the wage data beyond the initial wage of postunemployment jobs, we have added two pivotal terms to the equation: time-varying job tenure and its interaction with the dummy variable for sectoral choice. Thus, whereas existing studies explain how the duration of unemployment affects the initial wage rate of the postunemployment job, our specification emphasizes that the duration of unemployment entails long-term effects on wages beyond the initial survey week point of the postunemployment job. More importantly, our specification compares the wage effect of job change between movers and stayers in *both* the long-term (λ_3) and short-term (λ_1) perspectives.

Several econometrics issues are worth noting. First, there are three potentially endogenous regressors on the right-hand side of equation (4), which make OLS estimates inconsistent. They are duration of unemployment ($\text{UnempDuration}_{in}$), sectoral choice ($\text{INTER}_i$), and the interaction of the choice variable with the current tenure ($\text{INTER}_i \times \text{CurrentTenure}_{i,m,t}$).

The simultaneous determination of wages and unemployment duration arises from the standard job search theory, human capital arguments, and unobserved heterogeneity. For example, productive search theory predicts a positive association of duration and postunemployment wages [Stigler (1962); Lippman and McCall (1976); Mortensen (1986)]. A negative association of duration and postunemployment wages is observed when some workers are more productive than observationally equivalent others, so that the former experience shorter periods of unemployment and higher postunemployment wages than the latter. However, in addressing the simultaneity, existing studies have focused on the joint determination of the duration of unemployment and the initial wage offer on

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the postunemployment job [Kiefer and Neumann (1979); Addison and Portugal (1989)]. The current paper differs from the previous studies in that it explores the simultaneity of the duration of unemployment and the *entire* wage stream that workers receive from the postunemployment job. That unemployment has long-term effects on wages reflects the view that transferable (or general) human capital deteriorates as workers experience longer durations of unemployment.[7] The existence of lagged duration effects, which may arise when unemployment results in a loss of productive human capital, would also suggest that the duration of unemployment has a cumulative effect on wages. Of course, long or repeated unemployment may itself create lasting effects [Heckman and Borjas (1980)]. Knowing this, a worker may reduce his/her reservation wage with lengthening or repeated spells of unemployment. All these arguments support our specification (4) in that the duration of unemployment has a negative impact on all wages workers receive from the postunemployment job.[8] At the same time, we believe that not only the initial but also all subsequent wages (or the wage growth rate) matter in determining whether a worker recommences employment or not. Moreover, as movers experience longer durations of unemployment (see Table 1), even λ_1 and λ_3 will be inconsistently estimated by OLS.

The dummy variable of sectoral choice, INTER_i, may also be contemporaneously correlated with the error term, when the choice of switching sectors is endogenously made. This possibility was raised by Neal (1995), among others. With the sectoral choice being endogenous, not only INTER_i but also $\text{INTER}_i \times \text{CurrentTenure}_{it}$ is contemporaneously correlated with the error term. We are therefore left with three right hand–side endogenous regressors in equation (4).

To obtain consistent estimates of the coefficients in equation (4), the three endogenous regressors are instrumented by the following instrumental variables: the number of children at the time of job separation (KIDS_{i0}), a dummy variable that equals one if an unemployment spell begins in a recession period (RECESSION_i), the employment size of the industry the former job belongs to (IndustrySize_{io}), and the difference in the employment growth rate between the industry a respondent's former job belongs to and the rest (GrowthDiffer_{i0}). In addition, interaction terms of these four instrumental variables with the current tenure variable are used as additional instruments.

Although the validity of these instruments will be statistically tested, a brief discussion is in order. As noted by Addison and Portugal (1989) and Neal (1995), among others, the number of children is more or less reservation wage–specific and not directly related to productivity. If unemployed in recession, one experiences lower opportunity cost of job search, other things being constant, which also tends to lower his/her reservation wage. Therefore, the first two variables work directly as instruments for the duration of unemployment. We also follow Neal (1995) in using the industry employment size as an instrument for the sectoral choice: being unemployed from a large industry makes search cost in the same industry low. We extend his idea in two directions. First, we use the industry

TABLE 1. Comparison of inter- and intrasector mobility in terms of unemployment duration and postunemployment wages

	PSID			NLSY		
	Mean duration of unemployment (weeks)	Wages		Mean duration of unemployment (weeks)	Wages	
		Initial drop	Growth		Initial drop	Growth
Intrasector	8.68	−0.0251	0.0033	12.90	0.0387	0.0447
Intersector	18.08	−0.1593	0.0342	15.76	0.0169	0.0505

Note: The initial wage drop is calculated by first subtracting the logarithm of the last real wage of the preunemployment job from the logarithm of the first real wage of the postunemployment job and averaging them across individuals. The wage growth is computed by calculating year-to-year growth rates of individuals' real wages from postunemployment jobs and averaging them across years and persons.

employment size by state, which would be more relevant than the nationwide industry employment level when regional mobility costs existed. Second, more importantly, we consider as an additional instrument for sectoral choice the difference in the employment growth rate between the industry a respondent's former job belongs to and the other industries. If a respondent's former industry grows faster than the other industries, it becomes easier to find a job in the same industry, which creates more incentive to stay in the same industry. Conversely, if other industries grow faster than the respondent's former industry, switching sectors becomes more beneficial. This employment growth difference is also measured at the state level. We believe that this last variable is also qualified as an instrumental variable, as it does not have any direct effect on a person's postemployment wages.[9]

Instead of constructing structural equations for the endogenous regressors, we simply use reduced form equations for them. As wage growth, unemployment duration, and sectoral choice are jointly determined,[10] both duration and sectoral choice are instrumented by all four instrumental variables in the first regressions. In addition, if these four instrumental variables are valid for each of $\text{UnempDuration}_{in}$ and INTER_i, four interaction terms of these four instruments and the current tenure become natural instruments for the third endogenous regressor, $\text{INTER}_i \times \text{CurrentTenure}_{i,m,t}$. Put together, with three right hand–side endogenous regressors and with eight instrumental variables excluded from the wage equation, the structural parameters in equation (4) are overidentified. We employ the generalized method of moments (GMM) estimation to identify the structural parameters of equation (4).

The second econometrics issue is that, in equation (4), some explanatory variables such as unemployment duration do not vary across time within each individual. Specifically, different observations for the same individual may share some common component of variance that is specific to the person's spell of unemployment. In addition, different individuals in the same year may share some common component of variance that is specific to the year. In these cases, the error term in equation (4) will be positively correlated across different observations for the same individual or across different individuals within the same year. Neglecting these correlations would bias estimated standard errors downward. To obtain appropriate standard error estimates, we implement GMM cluster-robust estimation, which produces standard error estimates that are robust to arbitrary heteroskedasticity and intracluster correlation.[11]

Third, neglecting the period out of labor force, the length of difference in equation (4), which is nothing but the sum of current tenure and the duration of unemployment, is different across individuals and time. It was suspected that the variance of the differenced error term tends to be positively correlated with the length of difference, as $\text{Cov}(\varepsilon_{it} - \varepsilon_{i0})$ decreases in length. A modified Breusch–Pagan test, however, accepts the null hypothesis of homoskedasticity at any standard significance level.

3. EMPIRICAL FINDINGS

Table 1 compares benefits/costs between movers and stayers.[12] On average, movers experience longer durations of unemployment than stayers do. Our calculation based on the PSID sample [which replicates Starr-McCluer (1993)] reveals that the ratio of the average duration of unemployment for movers to stayers is approximately 2. The actual ratio is likely to be greater than 2, considering that right-censored spells are dropped in the calculation, their average duration is particularly long, and they are more likely to be movers. In addition, movers in general suffer from greater initial wage loss (or less wage gain) than stayers do, as shown in column (2).

Movers, however, enjoy higher wage growth in subsequent years on the post-unemployment job than stayers do. In the PSID sample, the average wage growth rate for movers is 3.4% per year, which is much higher than the 0.3% for stayers. These results are qualitatively confirmed by the NLSY data as well, as shown in columns (5)–(8). Although the above results are suggestive, these estimates may be contaminated by the differences between movers and stayers in productive attributes, which is the reason we turn to the formal regression analyses below.

3.1. OLS Results

Table 2 reports evidence from the PSID sample. All specifications except for column (5) are based on equation (4), where all observable and unobservable individual-specific but time-independent characteristics are "differenced out" by focusing on wage changes between two different time points. As another way of controlling for these fixed effects, column (5) uses the logarithm of the real wage level as the dependent variable and adds as an additional regressor the wage rate observed at the last survey week point of the preunemployment job, which, according to equation (2), contains information on individual-specific fixed effects, observable or not. Estimation results in the first five columns are based on the entire sample, whereas those in columns (6) and (7) are obtained by dividing them into two cases: (i) the sample of voluntary mobility and (ii) the sample of involuntary mobility.[13]

Focusing first on column (1), our basic model, the estimated coefficient of INTER (the dummy variable for intersectoral mobility) is not only statistically significant but also empirically important, implying that movers experience greater initial wage decline than stayers do, other things being equal. Similar evidence has been documented by a number of studies and interpreted as reflecting, among other things, the loss of sector-specific human capital associated with sectoral shifts. Specifically, the initial wage loss is 14% greater for movers than stayers. This figure is very similar to that reported by Addison and Portugal (1989, p. 282). Movers, however, enjoy higher wage growth in subsequent years than stayers do. The estimated coefficient of INTER $\times$ TENURE shows that the mover–stayer difference in the wage growth associated with an additional year of postunemployment tenure

TABLE 2. The effects of intersector movement on initial wages and wage growth of the postunemployment job: Estimates from OLS regression (evidence from the PSID sample)

	All workers					Voluntary	Involuntary
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
INTER	−0.1498***	−0.1465***	−0.1560***	−0.1493***	−0.1764***	−0.0749*	−0.1092***
	(0.0181)	(0.0181)	(0.0180)	(0.0270)	(0.0171)	(0.0455)	(0.0353)
TENURE	0.0083***	0.0121*	0.0091	0.0324***	0.0064***	0.0189**	0.0322***
	(0.0018)	(0.0063)	(0.0062)	(0.0068)	(0.0017)	(0.0080)	(0.0053)
TENURE_SQUARED	—	—	—	−0.0023***	—	—	—
				(0.0006)			
INTER × TENURE	0.0439***	0.0429***	0.0434***	0.00448***	0.0439***	0.0310**	0.0252***
	(0.0037)	(0.0037)	(0.0037)	(0.0139)	(0.0035)	(0.0109)	(0.0074)
INTER × TENURE_SQUARED	—	—	—	−0.0001	—	—	—
				(0.0013)			
Log(UnempDuration)	−0.0352***	−0.0356***	−0.0383***	−0.0356***	−0.0345***	−0.0443***	−0.0582***
	(0.0041)	(0.0042)	(0.0042)	(0.0041)	(0.0039)	(0.0137)	(0.0106)
TENURE_0	−0.0045***	−0.0045***	−0.0033***	−0.0045***	−0.0027***	−0.0048	−0.0118***
	(0.0007)	(0.0007)	(0.0007)	(0.0007)	(0.0007)	(0.0043)	(0.0020)
AGE	0.0136***	0.0123***	0.0126***	0.0126***	0.0192***	0.0338***	0.0152**
	(0.0033)	(0.0033)	(0.0033)	(0.0033)	(0.0031)	(0.0108)	(0.0074)
AGE_SQUARED	−0.0002***	−0.0002***	−0.0002***	−0.0002***	−0.0002***	−0.0004***	−0.0002**
	(0.00004)	(0.00004)	(0.00004)	(0.00004)	(0.00004)	(0.0001)	(0.0001)
SMSA	0.0328***	0.0357***	0.0204**	0.0342***	0.0552***	0.1046***	0.0328
	(0.0094)	(0.0095)	(0.0095)	(0.0094)	(0.0090)	(0.0322)	(0.0222)
INTERCEPT	−0.1765***	−0.1564**	−0.2234***	−0.1959***	0.1329**	−0.6304***	−0.1532
	(0.0645)	(0.0737)	(0.0740)	(0.0649)	(0.0627)	(0.1992)	(0.1419)

Table 2. Continued

	All workers					Voluntary	Involuntary
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Log(WAGE)_0	—	—	—	—	0.7810*** (0.0103)	—	—
YEAR DUMMIES	NO	YES	YES	NO	NO	NO	NO
OCCUPATION DUMMIES_0	NO	NO	YES	NO	NO	NO	NO
INDUSTRY DUMMIES_0	NO	NO	YES	NO	NO	NO	NO
R^2	.100	.108	.144	.105	.648	.147	.179
No. of year-person obs/spells			3,742/1,152			570/256	1,050/378

Note; In all specifications except for (5), the dependent variable is change in the logarithm of the real wage rate between year t of the current job and the ending year of the most recent job. In specification (5), it is the logarithm of the real wage rate of the current job in year t. X_0 is the value of variable X observed at the ending point of the most recent job. All the other variables pertain to the postunemployment job. INTER equals one for intersector movers, and zero for intrasector movers.

*, **, and ***Statistical significance at the 10%, 5%, and 1% level, respectively.

is 4.3%, which is also significant at any conventional significance level. Our calculation show that the 14% of the additional short-term wage loss associated with sectoral shifts is overturned in approximately four years by the higher wage growth of movers. That is, in four years on the postunemployment job, wages become higher for movers than for otherwise comparable stayers.

This finding is not significantly affected when year dummies are additionally included [column (2)] or when all sets of year, occupation (one-digit level), and industry (one-digit level) dummies are included [column (3)]. When squared tenure is additionally included for both inter- and intrasectoral mobility cases [column (4)], the real wage rate appears as a concave function of the current tenure for stayers, whereas it is a linear function for movers, which tends to make movers' long-term gains even greater. This finding remains valid even when we adopt a different approach of controlling for individual-specific fixed effects in column (5).[14]

In the last two columns, we split the sample between voluntary and involuntary cases and investigate any heterogeneous behavior between the two groups. Although the results are qualitatively very similar between the two groups, some differences arise in a quantitative sense. Compared with involuntary job changers, voluntary job changers experience smaller initial wage drops but greater wage growth when they cross industrial lines. (Readers should note that, in the current paper, both "movers" and "stayers" are job changers and are referred to "sector-movers" and "sector-stayers," respectively, whether or not they change jobs voluntarily.) Our calculation shows that it takes three years for movers' wages to catch up with stayers' wages when we confine our analysis to voluntary job changers. For involuntary job changers, the catch-up period turns out to be somewhat longer, five years. Despite the difference, the current results show that, even for involuntary job changers, the greater initial wage loss associated with sector switch tends to be compensated for by the higher wage growth for movers.[15]

In all columns of Table 2, the duration of unemployment reduces postunemployment wages significantly, implying that the negative wage effects such as human capital depreciation, stigma effects, and declining reservation wages associated with longer unemployment duration outweigh the positive effects of productive job search activities. However, the coefficient of the duration of unemployment may be inconsistently estimated by OLS due to the simultaneity of duration and postunemployment wages. Despite the likely bias, we include an interaction of the dummy for sectoral choice and the duration of unemployment in our basic model as an additional regressor, and find that the estimated coefficient of the interaction term is −0.0724, with corresponding standard error estimate 0.0104, implying that the adverse wage effect of unemployment experience is greater for movers than stayers, other things being constant. This exercise makes little difference in our main finding that the additional short-term wage loss associated with sectoral shifts is overturned in approximately four years by the higher wage growth of movers. For the rest of the paper, when we elaborate our arguments, we will revert to our preferred specification in column (1).

Table 3 presents evidence from the NLSY sample. We resort to our basic model in the first column of Table 2. Sector Classification 1 is used to obtain the results. Focusing on estimates in column (1), movers suffer from an additional initial wage loss of only 5.8%, in comparison to the 14% from the PSID sample. As in the PSID sample, however, the wage growth rate is higher for movers than stayers. Figures in column (1) show that it takes two years for movers' wages to catch up with stayers' wages. Estimates in column (2) and (3) show that, as in the PSID sample, the catch-up period is shorter among the voluntary group (only one year) than among the involuntary group (three years). These observations are generally preserved when we use the nonworking period instead of the unemployment period [columns from (4) to (6)]; when we include year dummies, occupation dummies, and industry dummies in the wage equation (not shown in the table); or when an alternative industry classification method is used (Table A.2).

3.2. Accounting for Endogeneity of Sectoral Choice and Duration of Unemployment

In testing the validity of the choice of instruments, we adopt Baum et al.'s (2003) strategy: at the minimum, (1) in the first-stage regression of each potentially endogenous regressor on the set of instruments, the null hypothesis that instruments are jointly insignificant is always rejected, and (2) the null hypothesis of the Hansen J (Sargan test in case standard two-stage least squares is applied) test of overidentifying restriction is never rejected.[16] Upon confirming validity of the set of selected instruments,[17] we conduct the Wu–Hausman F-test of the null hypothesis that the three potentially endogenous regressors are actually exogenous. In the PSID sample, the null hypothesis of exogeneity is rejected at any conventional level, as the F-test statistic is 3.88 with corresponding p-value .0088. Then we test if the resulting endogeneity hinges solely on the simultaneity of unemployment duration (or sectoral choice) and the initial wage rate received on the reemployed job, as often observed in the existing literature. To that effect, we drop initial wage observations from the sample and redo the same Wu–Hausman test. The F-value is 2.33 with associated p-value being .0721, implying that the duration of unemployment or sectoral choice is jointly determined not just with the initial wage rate but also with the subsequent wages received from the postunemployment job. This suggests that although the duration of unemployment has long-term effects on wages, possibly due to many factors, such as the loss of general human capital, scar effects, and declining reservation wages, job seekers also consider the entire wage stream they expect to receive from a new job when deciding whether to recommence employment or whether to cross the sectoral line.[18]

Accounting for the two types of endogeneity, however, does not change our main conclusion, as is evident in our GMM results in Table 4. The first three columns report estimated first-stage reduced form equations, whereas the last column reports the estimated wage equation.[19] Although not reported, in all first-stage regressions, tests of instrument relevance strongly rejected the null hypothesis that

TABLE 3. The effects of intersector movement on initial wages and wage growth of the postunemployment job: Estimates from OLS regression (evidence from the NLSY sample: Sector Classification 1)

	Unemployment spell			Nonworking spell		
	All	Voluntary	Involuntary	All	Voluntary	Involuntary
INTER	−0.0580**	−0.0144	−0.0953***	−0.0261	0.0191	−0.0991***
	(0.0228)	(0.0340)	(0.0304)	(0.0180)	(0.0229)	(0.0297)
TENURE	0.0385***	0.0527***	0.0316***	0.0351***	0.0388***	0.0285***
	(0.0043)	(0.0074)	(0.0051)	(0.0030)	(0.0038)	(0.0049)
INTER × TENURE	0.0382***	0.0322***	0.0376***	0.0295***	0.0265***	0.0332***
	(0.0053)	(0.0088)	(0.0066)	(0.0040)	(0.0050)	(0.0065)
Log(UnempDuration$n$)	−0.0406***	−0.0303***	−0.0338***	−0.0405***	−0.0381***	−0.0340***
	(0.0062)	(0.0094)	(0.0084)	(0.0041)	(0.0051)	(0.0075)
TENURE_0	−0.0360***	−0.0423***	−0.0292***	−0.0290***	−0.0296***	−0.0298***
	(0.0036)	(0.0058)	(0.0044)	(0.0028)	(0.0036)	(0.0045)
AGE	−0.0164	−0.0449**	0.0165	−0.0040	−0.0185	0.0213
	(0.0131)	(0.0193)	(0.0176)	(0.0108)	(0.0140)	(0.0171)
AGE_SQUARED	0.0002	0.0007**	−0.0004	−0.0000	0.0002	−0.0004
	(0.00002)	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0003)
SMSA	0.0508***	0.0610**	0.0157	0.0801***	0.0919***	0.0367
	(0.0175)	(0.0272)	(0.0224)	(0.0144)	(0.0190)	(0.0223)
INTERCEPT	0.4246**	0.8040***	−0.0595	0.2533	0.4675**	−0.1414
	(0.1961)	(0.2886)	(0.2636)	(0.1637)	(0.2132)	(0.2589)
R^2	.119	.132	.109	.091	.097	.081
No. of year-person obs/spells	5,397/2,477	2,720/1,237	2,608/1,122	9,033/3,964	5,625/2,463	3,228/1,421

Note: In all specifications, the dependent variable is change in the logarithm of the real wage rate between year t of the current job and the ending year of the most recent job. See the text for the definition of Sector Classification 1. Estimation results with Sector Classification 2 are presented in Table A.2. TENURE_0 is the job tenure observed at the ending point of the most recent job. All the other variables pertain to the postunemployment job. INTER equals one for intersector movers, and zero for intrasector movers. Inclusion of year dummies, industry dummies, and/or occupation dummies makes little difference in any of the above results.
*, **, and ***Statistical significance at the 10%, 5%, and 1% levels, respectively.

TABLE 4. Correcting for endogeneity of the duration of unemployment and sectoral choice: Estimates from generalized method of moments estimation

	First stage regressions			GMM estimation of wage change equation
	Log(Unemp Duration)	INTER	INTER × TENURE	
INTER	—	—	—	−0.0151
				(0.2091)
TENURE	−0.0037	0.0057	0.2431	0.0086*
	(0.01272)	(0.0045)	(0.0369)	(0.0051)
INTER × TENURE	—	—	—	0.0428*
				(0.0244)
Log(UnempDuration)	—	—	—	−0.1079**
				(0.0479)
TENURE_0	−0.0007	−0.0032***	−0.0161***	−0.0015**
	(0.0031)	(0.0009)	(0.0045)	(0.0007)
AGE	0.0228	−0.0202***	0.0047	0.0190***
	(0.0139)	(0.0053)	(0.0198)	(0.0069)
AGE_SQUARED	−0.0004***	0.0002***	−0.0003	−0.0003***
	(0.0002)	(0.0001)	(0.0002)	(0.0001)
SMSA	0.2173***	0.0637***	0.2679***	0.0120
	(0.0402)	(0.0134)	(0.0624)	(0.0118)
RECESSION	0.2226***	0.0526**	−0.0325	—
	(0.0623)	(0.0224)	(0.0863)	
KIDS	0.0001	−0.0133	0.0887***	—
	(0.0227)	(0.0084)	(0.0327)	
INDUSTRY_SIZE	−0.0036**	−0.0010	−0.0054**	—
	(0.0017)	(0.0007)	(0.0028)	
GROWH_DIFFER	−0.1127	0.0401	−0.2037	—
	(0.1723)	(0.0615)	(0.2526)	
RECESSION × TENURE	0.0252**	0.0007	0.0676**	—
	(0.0125)	(0.0043)	(0.0328)	
KIDS × TENURE	−0.0095**	−0.0021	−0.0527***	—
	(0.0043)	(0.0016)	(0.0115)	
INDUSTRY_SIZE × TENURE	0.0002	−0.0000	0.0004	—
	(0.0003)	(0.0001)	(0.0010)	
GROWTH_DIFF × TENURE	0.1171***	0.0376***	0.3094***	—
	(0.0370)	(0.0128)	(0.0983)	
INTERCEPT	1.1596***	0.6850***	0.2191	−0.1892
	(0.2820)	(1100)	(0.3835)	(0.1513)
No. of year-person obs/spells	3,482/983			
F-statistic/*P*-value	20.34/.000	27.51/.000	34.76/.000	12.36/.000
Hansen *J* statistic/*P*-value	—	—	—	0.742/.981

Note: The PSID sample is used. The three endogenous regressors are INTER, TENURE, and INTER × TENURE. The last row presents statistics of the Hansen *J* test of overidentifying restrictions. TENURE_0 is the job tenure observed at the ending point of the most recent job. All the other variables pertain to the postunemployment job. INTER equals one for intersector movers and zero for intrasector movers.
*, **, and ***Statistical significance at the 10%, 5%, and 1% levels, respectively.

instruments are jointly insignificant. Moreover, estimated coefficients of instruments in the first-stage regressions generally have correct signs. For example, taking the interaction effects into account, the number of children tends to reduce unemployment duration, and those who are unemployed during recession tend to experience greater duration and are more likely to switch sectors, other things being constant. The larger the respondent's former industry is, or the faster that industry grows relative to the other industries, the more likely that person is to stay in the same sector. Correcting for the simultaneity bias also increases the negative impact of unemployment duration on wages, compared to the OLS results. This result is consistent with that found by Addison and Portugal (1989, p. 292). Due to the large standard error estimate, the estimated coefficient of INTER_i turns out to be insignificant. However, our previous OLS estimate of mover–stayer difference in the wage growth rate (4.3%) still remains identical even in the GMM results.[20]

3.3. Evidence from Balanced Samples

Table 5 presents evidence based on balanced samples of the PSID and the NLSY. Individuals are included in each balanced sample only when they report wages along with other control variables for the first five consecutive years. Wages observed from the sixth year of the postunemployment job are also excluded for both movers and stayers. The primary purpose of this exercise is to make the most straightforward comparison of movers and stayers by holding the sample composition constant over time. At the same time, for persons who experience multiple spells of unemployment, requiring five years of wage observations on postunemployment jobs eliminates effectively all previous spells except for the most recent one, because, as previously mentioned, to be included in the sample, all sample spells must overlap the period between January 1980 and December 1982. As a result, each individual experiences only one unemployment spell in this balanced sample. This is potentially important because all the samples used in the above analysis allow for multiple spells per person, and the estimates may be biased in the presence of dependence across unemployment spells.[21]

This process dramatically reduces the sample sizes for both data sets, thereby reducing the statistical significance of the resulting estimates. However, all our previous findings are still preserved qualitatively even in the balanced sample. In particular, for both samples, wages of the postunemployment job grow much faster for movers than for stayers.

4. A QUANTITATIVE EXPLANATION

In Section 3, we found that although movers suffer from short-term costs such as longer duration of unemployment and larger initial wage decline, they enjoy benefits from higher wage growth in subsequent years on the postunemployment job. Still, we have not yet demonstrated that our empirical findings are also *quantitatively* consistent with an equilibrium explanation. On the basis of the estimates

TABLE 5. Evidence from five-year balanced panels

Variable	PSID	NLSY
INTER	−0.0662	0.0326
	(0.0433)	(0.0503)
TENURE	0.0157***	0.0302***
	(0.0059)	(0.0092)
INTER × TENURE	0.0289**	0.0371***
	(0.0128)	(0.0122)
Log(UnempDuration*n*)	−0.0497***	−0.0100
	(0.0065)	(0.0114)
TENURE_0	−0.0045***	−0.0347***
	(0.0012)	(0.0064)
AGE	0.0115*	−0.0316
	(0.0061)	(0.0299)
AGE_SQUARED	−0.0002***	0.0003
	(0.0001)	(0.0005)
SMSA	0.0467***	0.1064***
	(0.0152)	(0.0296)
INTERCEPT	−0.1175	0.7417*
	(0.1192)	(0.4480)
R^2	.087	.107
No. of observations/spells	1,220/244	1,340/268

Note: In all specifications, the dependent variable is change in the logarithm of the real wage rate between year t of the current job and the ending year of the most recent job. Individuals are included in the sample only when they report wages along with other control variables for the first five consecutive years. Wages observed from the sixth year of the postunemployment job are also excluded for both intersectoral movers and intrasectoral movers. TENURE_0 is the job tenure observed at the ending point of the most recent job. All the other variables pertain to the postunemployment job. INTER equals one for intersector movers, and zero for intrasector movers.
*, **, and ***Statistical significance at the 10%, 5%, and 1% levels, respectively.

obtained in Section 3, the current section addresses this issue by simply comparing lifetime utilities of movers and stayers.

To simplify our discussion, we assume that the duration of the postunemployment job is finite. This assumption is necessary because our estimates of wage growth call for an infinite value for the expected lifetime utilities if a reemployed worker receives a wage forever.[22] Then, the lifetime utilities of movers (V_M) and stayers (V_S) can be expressed as

$$V_M = d_M \cdot c + \beta^{d_M}\{w[1 - \text{drop}_M][1 + \beta(1 + g_M) + \beta^2(1 + g_M)^2 + \cdots + \beta^T(1 + g_M)^T + \beta^T V^*]\}$$

and

$$V_S = d_S \cdot c + \beta^{d_S}\{w[1 - \text{drop}_S][1 + \beta(1 + g_S) + \beta^2(1 + g_S)^2 + \cdots + \beta^T(1 + g_S)^T + \beta^T V^*]\},$$

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TABLE 6. Calibration

	Full sample	Short-duration jobs excluded
Preferences	$\beta = 0.96$	
Wages of preunemployment job	$w = 1$	
Unemployment compensation	$c = 0.5$	
Unemployment duration	$d_M = 0.347$ years (18.08 weeks), $d_S = 0.166$ years (8.68 weeks)	
Value of remaining lifetime utility	$V^* = 0$	
Initial wage drop	$\text{drop}_S = 2.51\%$, $\text{drop}_M = 16.51\%$	$\text{drop}_S = 2.51\%$, $\text{drop}_M = 15.32\%$
Wage growth	$g_M = 5.22\%$, $g_S = 0.83\%$	$g_M = 4.90\%$, $g_S = 0.71\%$
Results	Required duration: 8.1 years Actual duration: movers = 4.4 years, stayers = 5.6 years	Required duration: 7.7 years Actual duration: movers = stayers = 7.3 years

Note: The discount factor (β), the unemployment compensation (c), the last wage of the preunemployment job (w), the duration of the postunemployment job (T), and the value of future lifetime utilities as of when reemployment is terminated (V^*) are assumed to be the same for both movers (= intersectoral movers) and stayers (= intrasectoral movers). The required duration represents the duration of the postunemployment job that satisfies the equilibrium condition, the equality of the lifetime utilities of movers (V_M) and stayers (V_S).

respectively, where d_i, drop_i, and g_i are the duration of unemployment, initial wage drop, and wage growth rate of the postunemployment job for $i = M$ (movers) and $i = S$ (stayers), respectively. The discount factor (β), the unemployment compensation (c), the last wage of the preunemployment job (w), the duration of the postunemployment job (T), and the value of future lifetime utilities as of when reemployment is terminated (V^*) are assumed to be the same for both movers and stayers.

Table 6 reports calibration results. β is assumed to be 0.96, w is normalized to be 1, and c is assumed to be 0.5. On the basis of the estimates in Table 1, d_M and d_S are regarded as 0.347 and 0.166 years, respectively. V^* is assumed to be zero for both movers and stayers.[23] drop_S is assumed to be 2.51%, which is also obtained from Table 1. The rest of Table 6 has two different results, depending on the sample used. Figures in the first column are based on the OLS estimates from the full PSID sample (the first column of Table 2), whereas those in the second column, to be discussed subsequently, are obtained from the sample that excludes (from the full PSID sample) those who are separated from the postunemployment job in a year.

First, focusing on the full sample, drop_M is computed at 16.51%, and g_M and g_S are 5.22% and 0.83%, respectively. Then the duration of the postunemployment job is pinned down by substituting the estimated parameter values into the equilibrium

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condition, $V_M = V_S$. Our calculation suggests that for an equilibrium to be sustained, or for the intersectoral movement to be a rational choice, the duration of the postunemployment job, T, should be around 8.1 years.[24] A quantitative explanation of sectoral mobility then depends on the actual completed duration of the postunemployment job experienced by movers and stayers. The average completed duration in the PSID sample is 5.6 years for stayers and 4.4 years for movers. This implies that the entire mobility may not be quantitatively consistent with an equilibrium model.

However, when we exclude from the full sample those who are separated from the postunemployment job in a year, who constitute 23% of the entire people in the full PSID sample, the average completed job duration appears almost the same for both movers and stayers at 7.3 years (last row of the second column of Table 6).[25] With the remaining sample, we redo our previous regression analysis as in the first column of Table 2 and determine the additional initial wage loss associated with sector shifts to be 14% and the wage growth rates of movers and stayers to be 4.90% and 0.71%, respectively. All these figures are very precisely estimated. With these new estimates, the required duration of the postunemployment job in equilibrium become 7.7 years, implying that both types of labor mobility can be *quantitatively* consistent with an equilibrium framework.

Nevertheless, the question remains of why some workers who are separated from the postunemployment job in a year switch sectors, when they do not stay in the new sector long enough to enjoy the high wage growth rate. Our estimates suggest that inter- and intrasectoral movements of workers are quantitatively explained by an equilibrium framework only for a major group of workers who move with longer term perspectives.

5. SUMMARY

In this paper we attempt to explain why workers change sectors upon job separation, when intersectoral shifts are accompanied by the following costs. First, movers experience longer durations of unemployment than stayers do. Our calculations based on the PSID sample [which replicates Starr-McCluer (1993)] reveal that the ratio of average unemployment duration for changers to that for stayers is approximately 2. The actual ratio will be greater than 2 when one takes into account that the right-censored spells, which are more likely to be those of movers, are dropped in the calculation. Second, other things being equal, the initial wage loss is approximately 14% greater for movers than stayers.

Our answer lies in recognizing that movers enjoy higher wage growth in subsequent years of the postunemployment job than stayers do. The mover–stayer difference in the wage growth associated with an additional year of postunemployment tenure is estimated at 4.3%. These figures suggest that the additional short-term wage loss associated with sector change is overturned in four years by the greater wage growth of movers. The findings in the current study clearly demonstrate that the true economic costs of sectoral mobility tend to be overstated in existing studies. Calibration of a simple lifetime utility model demonstrates that movers and stayers can coexist in an equilibrium framework, at least for a major group of workers who move with longer term perspectives.

Our empirical evidence also shows that the duration of unemployment or sectoral choice is jointly determined not just with the initial wage rate but also with the subsequent wages received from the postunemployment job. Although the duration of unemployment has long-term effects on wages due to factors such as the loss of general human capital, scar effects, or declining reservation wages, job seekers also consider the entire wage stream they expect to receive in the future when deciding whether to recommence employment or not or whether to cross industrial lines or not.

NOTES

1. Lilien's empirical work (1982) is based on a measure of industrial employment composition that acts as a proxy of sectoral movements of workers. Other studies such as Loungani et al. (1990), Brainard and Cutler (1993), and Shin (1997) rely on a proxy of sectoral shocks, which are assumed to lead to sectoral movements of workers. These studies show that a significant portion of unemployment fluctuations are explained by sectoral movers.

2. See Loungani and Rogerson (1989), Starr-McCluer (1993), and Shin and Shin (2008), among others.

3. It is important to distinguish worker mobility from job reallocation. Whereas Davis and Haltiwanger (1990) find that intersectoral job reallocation is negligible, Shin and Shin (2008) find that a significant portion of workers do move across sectors. Shin and Shin (2008), however, also note that even at the worker level, the mobility is higher within sectors than across sectors.

4. In our sample, the NBER-dated recession periods are January 1980–July 1980 and July 1981–November 1982.

5. We are also grateful to Gary Solon for sharing the algorithm, which adopts a more conservative standard than Altonji and Shakotko (1987) for classifying a worker as an employer stayer. See Solon et al. (1994) for details.

6. In this derivation, the current tenure observed at the last survey week point of the most recent job is assumed to be completed; that is, $\text{CurrentTenure}_{i,(m-1),0} = \text{PastTenure}_{m-1}$.

7. Using a special data set that combines data on individual workers and their employers to obtain an independent measure of productivity, Hellerstein et al. (1999) find that, for prime-age workers (aged 35–54) and older workers (aged 55 and over), productivity and earnings rise at the same rate. This supports the theoretical prediction of the general human capital model.

8. In fact, in Addison and Portugal's accounting framework (p. 283), current wages are adversely affected by all previous unemployment duration spells, as in equation (1). Data limitations, however, make them focus on the short-term wage effect of the duration of unemployment. Kim and Shin (2006) empirically demonstrate that current wages are negatively affected not only by the most recent unemployment spell but also by all previous spells. From a dynamic point of view, their frameworks are qualitatively identical to our specification that unemployment has lasting effects on all wages received from the postunemployment job. Another strand of studies focuses on the long-term effects of job displacement on wages and earnings. For example, Topel (1990), Ruhm (1991), Jacobson et al. (1993), and Stevens (1997) consistently find that much of the effect of the displacement on earnings and wages is permanent.

9. To compute the industry employment size and the employment growth difference variables by state and by year, we use individual data with appropriate weights received from the Current Population Surveys.

10. The nature of joint determination of duration and sectoral choice is easily understood. As a person experiences a longer duration of unemployment, and consequently loses more of the sector-specific human capital accumulated in his or her former industry, the opportunity cost of switching sectors diminishes, which increases the likelihood of sector switch. At the same time, the decision to switch sectors from a declining to a growing sector in searching for long-term wage gains is made at various short-term costs such as longer duration of unemployment.

11. For a general discussion of a robust variance estimate that adjusts for within-cluster correlation, see Williams (2000) and Wooldridge (2002) among others. For a discussion of cluster-robust estimation within the GMM context, see Baum et al. (2003). Little change is made, however, when either "within-year" or "within-individual" clustering is considered.

12. Also, see Table A.1 for summary statistics of other variables used in the regression analysis.

13. Whereas an equilibrium approach considers every unemployment case as voluntary, we examine these two subsamples separately to investigate possible heterogeneity in the current issue. In the PSID sample, respondents are classified as involuntary job changers if they left their preunemployment jobs because their companies went out of business or they were laid off/fired. Voluntary job changers are those who quit their pre-unemployment jobs to look for or take another job, or for pregnancy or family reasons. Similar classifications are applied to the NLSY sample.

14. The dummy variable SMSA equals one if the respondent lives in a small metropolitan statistical area. The dummy variable observed at the ending point of the most recent job (SMSA_0) is excluded from the equation because it appears insignificant in every specification. AGE_0 is also excluded from the equation to avoid its multicollinearity with the current age, the current tenure, and the duration of unemployment.

15. It should be noted that unemployment spells in the two subsamples summed up to 634, which is far less than the total number of spells, 1,152. This difference is mostly due to our exclusion from the subsamples of temporary layoffs, which are already included in the total sample. When these temporary layoffs are included in the subsample of involuntary job changers, the number of involuntary spells increases to 783, with the number of wage observations being 2,805. Our OLS results from this extended involuntary subsample produce estimated coefficients of INTER_i, $\mathrm{CurrentTenure}_{i,m,t}$, and $\mathrm{INTER}_i \times \mathrm{CurrentTenure}_{i,m,t}$ of -0.1838, 0.0069, and 0.0478, respectively, with respective standard error estimates of 0.0216, 0.0020, and 0.0043. Now, it takes four years for movers' wages to catch up with stayers' wages.

16. The full set of test results is available from the authors upon request. The statistics of the Hansen J test are reported at the bottom of Table 4.

17. We also test and find that marital status and the amount of nonwage income at the time of leaving the former job cannot be valid instruments, as they cannot be excluded from the wage equation.

18. We cannot conduct the test of endogeneity of regressors using our NLSY sample, as the explanatory power of instrumental variables is generally weak in the first stage regression.

19. Following an anonymous referee's comment, we do not exploit censored duration in obtaining the first-stage predicted value of the duration of unemployment, due to the possible inconsistency of estimated coefficients that may arise from nonlinearity of instrumental variables. For a similar reason, a linear probability model is used to predict the first-stage sectoral choice variable.

20. In our model, the total number of excluded instrument variables is eight. As noted by Hayashi (2000), among others, the Hansen–Sargan type of test of overidentifying restriction may have very little test power when we deal with a model with a large number of excluded instruments. As also found by Hahn and Hausman (2002), the size of bias associated with instrumental variable estimation increases in the number of instruments. To avoid the potential bias associated with using too many instruments, we try various subsets of the eight instruments based on a formal test (often called the *C* statistic or difference-in-Sargan statistic) and find results that are generally consistent with those OLS estimates in Table 2.

21. Although Heckman and Borjas (1980), Ellwood (1982, p. 350), Ruhm (1991, p. 322), and Choi and Shin (2002) find no evidence of occurrence dependence in unemployment spells in the United States labor market, Corak (1993) reports positive occurrence dependence across unemployment insurance spells in the Canadian labor market.

22. Because the estimated annual wage growth rate (5.2%, first column in Table 2) of movers is higher than the discount rate in a reasonable range (around 4%), the infinite sum of the discounted wages produces an infinite value.

23. Our results are quite robust with respect to variation of parameter values. For example, when we vary the discount factor between 0.96 and 0.99 and V^* between 0 and 10, the main results hardly changes.

24. The required duration will be reduced considering that, although the wage growth rate is constant for movers, it decreases in tenure for stayers [column (4) of Table 2].

25. This justifies our simplifying assumption of equating the duration of the postunemployment job for movers with that for stayers in calibrating parameters.

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APPENDIX

TABLE A.1. Summary statistics

	PSID		NLSY (not working)		NLSY (unemployment)	
INTER	0.2934***		0.6726***		0.7062***	
	(0.0134)		(0.0075)		(0.0092)	
UNEMP DURATION	11.664***		33.305***		14.685***	
(in weeks)	(0.448)		(0.9216)		(0.3622)	
ln	2.0397***		1.6179***		1.6049***	
(REAL WAGE_0)	(0.0136)		(0.089)		(0.0108)	
TENURE_0	3.6474***		1.7345***		1.6141***	
(in years)	(0.1724)		(0.0355)		(0.0419)	
EDUCATION	11.629***		12.49***		12.305***	
	(0.066)		(0.0359)		(0.0427)	
MALE	0.8123***		0.5492***		0.6082***	
	(0.0116)		(0.0079)		(0.0098)	
WHITE	0.5955***		0.6438***		0.6182***	
	(0.0145)		(0.0076)		(0.0098)	
YEAR		84.808***		90.734***		90.306***
		(0.0507)		(0.0515)		(0.066)
ln		2.0589***		1.789***		1.7433***
(REAL WAGE)		(0.0072)		(0.006)		(0.0071)
TENURE		3.7002***		3.0194***		2.7904***
(in years)		(0.0477)		(0.0333)		(0.0401)
AGE		37.229***		29.672***		29.214***
		(0.1777)		(0.0555)		(0.0715)
SMSA		0.5531***		0.7684***		0.7568***
		(0.008)		(0.0044)		(0.0058)
No. of observations	1,151	3,742	3,964	9,033	2,477	5,397

Note: Consumer Price Index (100 for 1982–1984) is used to derive real wages. In each column, figures flushed to the left and the right pertain to spells and year–person observations, respectively. *X*_0 is the value of variable *X* observed at the ending point of the most recent job. INTER equals one for intersector movers, and zero for intrasector movers. For the NLSY sample, Sector Classification 1 is adopted to compute the proportion of intersectoral movers among the total movers.

*** Statistical significance at the 1% level.

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TABLE A.2. The effects of intersector movement on initial wages and wage growth of the postunemployment job: Estimates from OLS regression (evidence from the NLSY sample: Sector Classification 2)

	Unemployment spell			Nonworking spell		
	All	Voluntary	Involuntary	All	Voluntary	Involuntary
INTER	−0.0245	−0.0188	0.0022	0.0036	0.0185	−0.0089
	(0.0323)	(0.0462)	(0.0453)	(0.0258)	(0.0318)	(0.0430)
TENURE	0.0320***	0.0641***	0.0159***	0.0240***	0.0280***	0.0166***
	(0.0041)	(0.0074)	(0.0050)	(0.0031)	(0.0039)	(0.0048)
INTER × TENURE	0.0312***	0.0311***	0.0139*	0.0249***	0.0237***	0.0215***
	(0.0063)	(0.0097)	(0.0085)	(0.0049)	(0.0061)	(0.0082)
Log(UnempDuration)	−0.0062	0.0304**	−0.0347***	−0.0467***	−0.0460***	−0.0507***
	(0.0088)	(0.0136)	(0.0126)	(0.0058)	(0.0068)	(0.0116)
TENURE_0	−0.0317***	−0.0227***	−0.0305***	−0.0284***	−0.0256***	−0.0302***
	(0.0043)	(0.0068)	(0.0057)	(0.0034)	(0.0043)	(0.0055)
AGE	−0.0070	−0.0224	0.0149	−0.0204	−0.0532**	0.0293
	(0.0199)	(0.0295)	(0.0267)	(0.0169)	(0.0215)	(0.0263)
AGE_SQUARED	−0.0000	−0.0000	−0.0002	0.0003	0.0008**	−0.0004
	(0.0003)	(0.0005)	(0.0004)	(0.0003)	(0.0003)	(0.0004)
SMSA	0.0651***	0.0907**	0.0537	0.0293	0.0265	0.0432
	(0.0246)	(0.0366)	(0.0329)	(0.0195)	(0.0246)	(0.0311)
INTERCEPT	0.2697	0.5584	−0.1362	0.5627	1.0649***	−0.2419
	(0.3046)	(0.4527)	(0.4070)	(0.2596)	(0.3320)	(0.4010)
R^2	.099	.176	.055	.075	.085	.062
No. of year-person obs/spells	2,492/645	1,222/341	1,241/294	4,340/1,139	2,743/747	1,525/368

Note: In all specifications, the dependent variable is change in the logarithm of the real wage rate between year t of the current job and the ending year of the most recent job. See the text for the definition of Classification 2. TENURE_0 is the job tenure observed at the ending point of the most recent job. All the other variables pertain to the postunemployment job. INTER equals one for intersector movers, and zero for intrasector movers. Inclusion of year dummies, industry dummies, and/or occupation dummies makes little difference in any of the above results.

*, **, and *** Statistical significance at the 10%, 5%, and 1% levels, respectively.