WITHIN- AND CROSS-FIRM MOBILITY AND EARNINGS GROWTH

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A widely accepted premise is that promotions within firms and mobility across firms lead to significant earnings progression. Existing research generally has examined cross-firm mobility separately from hierarchical advancement. Yet, as the authors' descriptive evidence from Danish panel data shows, how the two types of mobility interact is important for understanding earnings growth. Cross-firm moves at the nonexecutive level provide sizable short-run earnings growth (similar to the effect of being promoted to an executive position). These gains, however, appear modest compared with the persistent impact on earnings growth of promotions (either within or across firms) and subsequent mobility at a higher hierarchy level.

From the job search literature, we know that mobility across firms is an important contributor to the growth in wages that employees experience over their career (for a review, see Rogerson, Shimer, and Wright 2005). The personnel economics literature documents the importance of promotions for earnings progression (for a review, see Gibbons and Waldman 1999a). Yet existing theoretical and empirical research has generally examined cross-firm mobility separately from hierarchical advancement. As a result of the dichotomy in the literature, little evidence is currently available on the importance of interactions between within-firm and cross-firm mobility. This article presents descriptive evidence on the effects of within- and between-firm mobility on earnings growth using a rich Danish panel

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data set that provides information both on employer–employee matches and on broad hierarchy levels.

The impact of cross-firm mobility on earnings and earnings growth is the subject of a substantial literature. In an influential study, Topel and Ward (1992) found that the wage increases that employees experience when moving to new employers account for more than one-third of the wage growth during the first decade of the working life of white men in the United States.[1] Using a structural estimation of an equilibrium search model with Danish data similar to ours, Bagger, Fontaine, Postel-Vinay, and Robin (2014) concluded that job searches explain a large part of the wage growth early in a career and that it drives most of the wage dynamics afterward. The mechanism leading to earnings growth in this model is similar to the one in Postel-Vinay and Robin (2002), in which firms do not commit to long-term wage growth but, instead, renegotiate the worker's piece rate each time he or she obtains an attractive outside offer. Bargaining ends with the firm that values the worker most (i.e., the more productive firm) retaining or poaching the worker, and in this process, the worker captures part of the surplus relative to the less productive match. As a result, workers matched with more productive firms experience higher wage growth on average than those matched with less productive firms. Another interesting model is presented in Carrillo-Tudela and Kaas (2011). They modeled a frictional labor market with asymmetric information about workers' abilities. In equilibrium, firms offer contingent contracts that lead to wage increases for high-ability types and wage decreases for low-ability types. Nevertheless, although these models capture the effect of job searches on wage growth both within and between firms, they do not explicitly account for hierarchy levels in firms.

The role of position changes within firms is the subject of a different literature, mostly based on data covering individual firms or particular occupations. In two influential articles, Baker, Gibbs, and Holmström (1994a, 1994b) empirically studied the internal workings of a firm and evaluated to what extent the main theories of earnings dynamics, such as on-the-job training, learning, and incentives, explained their findings. One particularly important empirical finding established in this research is that promotions are an important source of earnings growth and that immediate wage increases at promotion account for only part of the average wage difference across hierarchy levels. This often-replicated finding suggests that much of the gain from promotions comes in the form of faster compensation growth at higher levels in the hierarchy.[2]

To our knowledge, only two studies have estimated the effects on earnings of both within-firm hierarchical advancement and between-firm

[1]Other contributions are, for example, Bartel and Borjas (1981); Antel (1986, 1991); Mincer (1986); Altonji and Shakotko (1987); Topel (1991); Keith and McWilliams (1999); Altonji and Williams (2005); Dustmann and Meghir (2005); Buchinsky, Fougere, Kramarz, and Tchernis (2010).

[2]Other contributions are, for example, Medoff and Abraham (1980, 1981); Lazear (1992); Chiappori, Salanie, and Valentin (1999); Seltzer and Merrett (2000); Treble, Van Gameren, Bridges, and Barmby (2001); Booth, Francesconi, and Frank (2003); Dohmen, Kriechel, and Pfann (2004); Gibbs and Hendricks (2004); Belzil and Bognanno (2008).

mobility. McCue (1996) computed from average real wage changes in the Panel Study of Income Dynamics (PSID) that around 10% of the wage growth that an individual experiences over the first decade in the labor market can be attributed to promotions. Around 24% of the 10-year wage growth is linked to cross-firm moves. Da Silva and van der Klaauw (2011) used Portuguese matched employer–employee data and controlled for individual unobserved heterogeneity. They found substantial returns to promotions and cross-firm transitions, each of which provided an immediate wage increase of around 5%.

Our results show that it is important to allow for the additional details that some cross-firm moves are, in fact, promotions (or demotions) and that lateral moves occur at different hierarchy levels. Paying attention to these details, while at the same time accounting for unobserved individual heterogeneity, helps clarify the relative contributions that different types of between-job mobility make to the long-run earnings progression. In our sample of men with a stable labor-force attachment, cross-firm moves provide sizable short-run gains. Switching employers at the nonexecutive level (which constitutes more than 90% of our sample) is comparable to receiving a within-firm promotion to an executive-level job. The one-off gain from a cross-firm move, however, is relatively modest in comparison to the persistent impact that promotions, either within or across firms, and subsequent mobility at the executive level have on earnings growth. Using our estimates to compute the 10-year log growth rates for a university graduate who switches employers or receives a promotion early on in his or her career, we find that 12 to 18% of the total growth can be attributed to the promotions. Only 2 to 9% stem from the earnings gain that the employee experiences when switching employers. A final important result is that the returns to vertical transitions across firms exhibit an interaction effect: the short-run gain from a cross-firm promotion exceeds the sum of the premia of a (within-firm) promotion and a (lateral) cross-firm move, and this gain is only partly reversed for a cross-firm demotion.

To establish the role that interactions between cross-firm mobility and hierarchical transitions play for earnings growth, we use register-based linked employer–employee data from Denmark. Denmark has a flexible labor market with high cross-firm mobility, and in these respects, it is similar to the United Kingdom and the United States (Jolivet, Postel-Vinay, and Robin 2006). Our data allow us to trace employee mobility both within and between firms and provide information on compensation as well as on a large set of background variables. In the spirit of the earnings dynamics literature,[3] our estimation procedure pays attention to the importance of permanent and transitory shocks to the earnings process. Our approach allows for predetermined mobility (i.e., mobility that is driven by past earnings), and it accounts for the potential fixed effects in earnings growth.

[3]For example, Lillard and Willis (1978); Lillard and Weiss (1979); MaCurdy (1982); Abowd and Card (1989); Baker (1997); Meghir and Pistaferri (2004); Altonji, Smith, and Vidangos (2009); Browning, Ejrnaes, and Alvarez (2010).

Our main results are based on a sample of employees with stable careers in the private sector. Even though a gender wage gap is apparent, the relative returns to mobility are remarkably similar for men and women. We may interpret this to mean that men and women who are solidly attached to the labor market face similar incentives to look for alternative employment or to compete for a promotion. But the likelihood of a cross-firm move or a promotion, which are the two types of flows associated with sizable earnings gains, is lower for women. This implies that between-firm mobility and hierarchical transitions, despite similar relative returns for men and women conditional on a move, tend to increase gender differences in earnings. We also estimate the model separately for different education groups and find that the returns to mobility increase with education level.

Data and Descriptive Statistics

Our study uses register-based information on all establishments and residents in Denmark from Statistics Denmark's Integrated Database for Labor Market Research (Integreret Database for Arbejdsmarkedsforskning, IDA). The database provides detailed information on mobility across firms; unique identifiers allow us to follow individuals and establishments over time (matches are recorded once a year in November). Further, the data permit us to construct a measure of hierarchical placement. Using the first digit of the Danish International Standard Classification of Occupations (DISCO) codes, we can distinguish "executives" (employees who manage organizations or departments; major group 1, comprising corporate managers and general managers), from "nonexecutives" (subsuming all other major groups).[4]

Hierarchical Placement

For the purposes of studying cross-firm moves we particularly need to have a reliable measure of hierarchical placement that is comparable across firms.[5] Our hierarchical placement variable has two advantages.

[4]The Ministerial Order on the Statistics Denmark Act (Bekendtgørelse af lov om Danmarks Statistik) requires every employer in Denmark to report annually an occupational classification code for each of its full-time employees. Reporting to Statistics Denmark normally takes place directly through the company's electronic salary systems, using DISCO codes. These are the Danish version of the International Standard Classification of Occupations by the International Labour Organization (ILO). For documentation, see http://www.ilo.org/public/english/bureau/stat/isco/.

[5]Studies on within-firm promotions can often draw on a firm's personnel records for hierarchical classification (e.g., Baker et al. 1994a, 1994b). Occupational classifications are sometimes used, but they often are very imprecise measures of hierarchical placement unless the classifications are specifically designed to capture hierarchy (rather than job-related skills). For example, Hunnes (2012) relied on hierarchical classifications drawn up by the Confederation of Norwegian Enterprise for wage bargaining. Classifications based on survey data have problems distinguishing mere wage increases from promotions involving real shifts in jobs. In our case, the occupational classifications of executives and nonexecutives are precise measures of hierarchy. Smith, Smith, and Verner (2013) used Danish data similar to ours. To obtain finer-grained hierarchical placements within the executive rungs, they merged the data with information from the private Danish data account register Experian and used earnings data to sort out the misclassifications. This allowed them to also classify employees as chief executive officers (CEOs; with the additional requirement that the annual income be among the top 10), vice presidents (among the top 25 earners), and executives (just below these ranks).

First, it provides a clean measure of a promotion that involves an actual change in position.[6] Such a shift in the employee's production technology is central to prominent theoretical models of wage and promotion dynamics (e.g., Bernhardt 1995; Gibbons and Waldman 1999b, 2006). Thus, our concept of promotion corresponds well with theoretical models in which jobs are assigned based on employees' comparative advantage. Those promoted are given jobs in which the output is more sensitive to ability and accumulated human capital than those who are not promoted. This clearly distinguishes promotions from mere wage increases.

Second, our measure has a consistent interpretation across the wide spectrum of firms covered by our data. This helps us avoid some of the problems encountered with promotion measures based on organizational charts, occupational classifications, and self-reports from employees or employers. Their firm- or industry-specific nature complicates comparisons. First, members of an organization often do not perceive as a promotion what the classification identifies as a change in hierarchical level. Da Silva and van der Klaauw (2011), for instance, reported that more than 70% of all moves classified as a change in hierarchical level in their data are not considered to be a promotion by the employer. Second, self-reported promotions that involve no position change are hard to distinguish from other elements of pay for performance. For example, in Pergamit and Veum (1999) and da Silva and van der Klaauw (2011), 40 to 50% of self- or employer-reported promotions involved no change in job description.[7]

One caveat is that we might underestimate the returns to promotions. Because we can only pick up moves from the nonexecutive to the executive level, some of the individuals classified as staying at the same hierarchy level may have, in fact, really received some sort of small promotion.

Sample Selection

Our aim is to shed light on how mobility affects earnings growth for those with stable careers. Therefore we analyze earnings patterns for core employees in private-sector establishments. Specifically, we follow employees who were continuously in full-time employment between 1994 and 2005 in private-sector establishments. Because small firms may lack internal labor markets, we restrict the sample to firms with at least 25 employees. To abstract from decisions to acquire further formal education or decisions about retirement, we further restrict the sample to employees who were between 30 and 45 years of age at the start of the panel in 1994.[8] Our earnings measure

[6]For a detailed discussion of the scope and reliability of the DISCO codes as measures of career mobility, see Frederiksen and Kato (2011).

[7]Pergamit and Veum (1999) exploited questions about the receipt of promotions in the 1990 wave of the National Longitudinal Survey of Youth. Da Silva and van der Klaauw (2011) used employer-reported promotions in the Portuguese *Quadros de Pessoal*.

[8]To obtain a sample of individuals who were continuously in full-time employment, we use the monthly social security record of employment status. Employer–employee matches are recorded only once a year, in November. Therefore, shorter employment periods (and associated flows), for instance, lasting from March to September of a particular year, cannot be picked up using our data. Given our focus on core employees with continuous employment histories, however, this does not seem problematic for our purposes.

Table 1. Descriptive Statistics, 1994

	Men	*Women*
Age (years, mean)	37.63	38.03
	(4.54)	(4.49)
Education (%)		
9 years (less than high school)	17.27	22.10
12 years (high school)	55.78	60.40
15 years (bachelor's degree or post-secondary professional training)	19.37	13.66
17 years (master's degree or higher)	7.60	3.85
Real labor income (DKK)[a]	345,320	262,914
	(123,289)	(82,147)
Firm size (number of employees. employee weighted average)	2,192	3,503
	(3,537)	(4,565)
Number of unique individuals	57,010	26,352
Person-year observations (1994 to 2005)	684,120	316,224

Notes: The table contains descriptive statistics for a balanced panel of employees who were ages 30 to 45 in 1994. The employees were continuously employed between 1994 and 2005 in private firms with at least 25 employees. Values in parentheses are standard deviations. DKK, Danish kroner.
[a]DKK 100 = US$12 (at year-2000 levels).

is annual labor income converted to year-2000 prices using Statistics Denmark's consumer price index. It comprises both base-pay components and variable-pay components, such as bonuses. The registers do not allow us to distinguish between these components.

We analyze separate samples for men and women, each giving a balanced panel in which every individual has a complete 12-year employment history. Table 1 presents the descriptive statistics. The male sample consists of 57,010 unique individuals with 684,121 person-year observations, and the female sample consists of 26,352 unique individuals with 316,224 person-year observations. At the start of the panel, the average employee in both samples is about 38 years old. Some noticeable gender differences emerge: men work in smaller firms than women, tend to be more educated than women, and earn about 30% more than women. The average man earns DKK 345,320 (Danish kroner in year-2000 prices, corresponding to around US$41,000); the average woman earns DKK 262,914 (around US$31,000).

Patterns of Mobility

The flows in this article are based on a comparison for each person of his or her primary employment relationships in November of consecutive years, when employer–employee matches were recorded. Table 2 shows the patterns for all the eight different types of cross-firm and within-firm moves.

In the male sample, 92% of all employees are in the nonexecutive layer. Most remain at that level, with 88% of them staying with the same firm (Stayer) and 11% moving laterally between firms (CF). A bit more than 1%

Table 2. Mobility Patterns, Probability of Annual Transition, 1994–2005

	Men[a]		*Women*[b]	
	Percentage	*Transition probability*[c]	*Percentage*	*Transition probability*[c]
A. All transitions	100	100	100	100
Within-firm moves		89.14		89.93
Cross-firm moves		10.86		10.07
B. Nonexecutive level	92.20	100	97.62	100
$\text{Nonexecutive}_{t-1} \rightarrow \text{Nonexecutive}_t$				
No move (Stayer)	80.97	87.82	87.37	89.50
Lateral move, cross-firm (CF)	9.80	10.63	9.75	9.99
$\text{Nonexecutive}_{t-1} \rightarrow \text{Executive}_t$				
Promotion, within-firm (PWF)	1.17	1.27	0.41	0.42
Promotion, cross-firm (PCF)	0.27	0.29	0.09	0.09
C. Executive level	7.80	100	2.38	100
$\text{Executive}_{t-1} \rightarrow \text{Executive}_t$				
No move (ExecStayer)	6.11	78.41	1.83	77.00
Lateral move, cross-firm (ExecCF)	0.57	7.33	0.16	6.62
$\text{Executive}_{t-1} \rightarrow \text{Nonexecutive}_t$				
Demotion, within-firm (DWF)	0.89	11.45	0.31	13.18
Demotion, cross-firm (DCF)	0.22	2.81	0.08	3.20

[a]Men: 627,110 person-year observations on transitions in 1995–2005 ($N = 57{,}010$ unique individuals).
[b]Women: 289,872 person-year observations on transitions in 1995–2005 ($N = 26{,}352$ unique individuals).
[c]Annual transition probability (percentage of group).

are promoted to the executive layer within the same firm (PWF); promotions across firm boundaries (PCF) account for 0.3%. Similarly, executives (who make up 8% of male employees) typically remain in that level; but less persistence is evident than for nonexecutives: 78% stay with the same firm (ExecStayer) and 7% move laterally across firms (ExecCF). Just slightly more than 11% of executives are demoted to nonexecutive positions within the firm (DWF), and slightly less than 3% cross firm boundaries and continue at the nonexecutive level (DCF).

Downward moves, hence, are not uncommon; but promotions (both within- and cross-firm) are about 1.3 times more frequent than demotions. Our data thus add to a number of studies showing that demotions are by no means exceptional; these include Hunnes (2012), a study of white-collar workers in Norway (with a promotion/demotion ratio of 2.67 for technical and 1.51 for administrative workers); Belzil and Bognanno (2008), a study of U.S. executives (promotion/demotion ratio of 1.1 or 5.1, depending on the definition of hierarchical levels); Frederiksen, Lange, and Kriechel (2012), a benchmark study of six private European and U.S.-based companies (promotion/demotion ratios ranging up to 2.0); Lluis (2005), an analysis of data from the German Socio-Economic panel (GSOEP) (promotion/demotion ratio of 1.1 after a wage-growth-based reclassification); Hamilton and MacKinnon (2001), a study of the Canadian Pacific Railway (promotion/demotion ratio 1.7); and Seltzer and Merrett (2000), a study of the 19th-century Union Bank of Australia (promotion/demotion ratio 2.1).

Keeping in mind that we look at a rather large promotion step from the nonexecutive to the executive level, our annual promotion rate of 1.6% tends to be lower than those in other studies, for example, 5.3 to 6.7% for Norwegian white-collar workers (Hunnes 2012), 3.8% in the GSOEP (Lluis 2005), 2.4 to 16% in the benchmark study by Frederiksen et al. (2012), 8% in a British financial-sector firm (Treble et al. 2001), 14.5% in a U.S. financial-sector firm (Baker et al. 1994a, 1994b), 7.2% for a large U.S. corporation (Gibbs and Hendricks 2004), and 20% for the Canadian Pacific Railway (Hamilton and MacKinnon 2001). Within a firm operating plants in two countries, Grund (2005) found a promotion rate of 1.2% in the German plant and of 8.4% in the U.S. plant.

Women are slightly less likely than men to make cross-firm moves (10% compared to 11% for men). Only 2.4% of women are employed at the executive level (compared to 7.8% for men). This is, in part, explained by a lower probability of promotion for women (0.5% compared to 1.6% for men) and a higher probability of demotion (16% compared to 14% for men).

Overall, the cross-firm mobility patterns are similar to those reported by McCue (1996) for the United States (men 11% and women 12%; using the PSID, 1976 to 1988) and higher than those reported by Lluis (2005) for Germany (6% for both men and women; GSOEP, 1985 to 1996). Indeed, in a cross-country comparison by Jolivet et al. (2006) based on the European Community Household Panel (1994 to 2001) and the PSID (1993 to 1996), Denmark belonged to the group with high job-to-job transition rates (15 to 20% in a three-year window), along with Ireland, the United Kingdom, and the United States. The middle range of job-to-job transition rates is found in Germany and the Netherlands, whereas rates well below 10% are found in Belgium, France, Italy, Portugal, and Spain.[9]

In the study by da Silva and van der Klaauw (2011) using Portuguese data, fewer than 20% of those in the sample have more than nine years of schooling and their average annual earnings are approximately US$9,000 (in year-2000 prices). In comparison, in our sample around 80% have more than the nine years of compulsory schooling and the average annual earnings are around US$30,000 (in year-2000 prices). Our data thus come from a more flexible labor market with a much more highly educated labor force and higher income levels, which in these respects is more similar to the labor markets in the United Kingdom and the United States.

Preliminaries

In the matter of data description, let us start with an ordinary least squares (OLS) estimation of the following earnings growth model:

[9]For further details on job-to-job mobility in Denmark see Frederiksen (2008), and for a general discussion of institutional features and issues related to mobility and earnings, see Eriksson and Westergaard-Nielsen (2009).

$$(1) \qquad \Delta \ln\left(Y_{it}\right) = \ln\left(Y_{it}\right) - \ln\left(Y_{i(t-1)}\right) = \alpha + \sum_{j=1}^{J} \mu_j M_{jit} + X'_{it}\beta + u_{it}$$

where $Y_{i,t}$ are the earnings data for individuals, $i = \{1,\ldots,I\}$ at dates $t = \{1,\ldots,T\}$. The right-hand side of Equation (1) consists of a constant, J mobility dummies (M_{jit}), a vector of controls (X_{it}), and a residual (u_{it}). The mobility dummies correspond to the flows CF, PWF, PCF, ExecStayer, ExecCF, DWF, and DCF (previously presented). The reference group is Stayer, nonexecutive employees who stay at that level in the same firm. Our control variables include a quadratic in age as well as education, sector, and year fixed effects.

Table 3 contains the estimation results. Moving to a new employer is associated with about 1% higher labor earnings growth for both men and women, as reported in columns (1) and (4). Columns (2) and (5) consider hierarchical transitions on their own. An upward move accelerates earnings growth by around 1 to 2%, whereas a downward move has no significant effect.

Interaction effects between cross-firm moves and hierarchical transitions are shown in Table 3, columns (3) and (6), respectively. Switching firms at the nonexecutive level (CF) yields around 1% higher growth relative to staying with the same employer at that level. A within-firm promotion yields roughly the same coefficient as CF for men and women. The biggest return is for a cross-firm upward move, with around 6% higher growth for men and 4% for women. Our estimates suggest that executive-level jobs are associated with a steeper earnings profile: earnings increase 0.6% faster than for nonexecutive Stayers. Furthermore, cross-firm mobility pays off more at the executive level, yielding around 3% higher growth for men and around 2% for women relative to ExecStayers. Finally, demotions appear to reset the growth of an executive's earnings at the rate of a nonexecutive Stayer. So, even though demotees do not suffer negative earnings growth in the year of their demotion (or only slight decreases in real wages),[10] they do lose out on the higher pay progression they would have enjoyed if they had remained executives.

Overall, the first impression is that lateral cross-firm mobility seems to count roughly as much as a within-firm promotion and that moving to a new firm tends to enhance the returns to vertical mobility. Note, however, that this model does not properly account for the covariance structure of earnings growth. Nevertheless, this exercise provides a descriptive backdrop that will help us to understand the role that transitory shocks play in the relationship between mobility and earnings progression in our econometric strategy. We return to this issue in the next section.

The Econometric Strategy

Following previous contributions to the earnings dynamics literature, our estimation procedure accounts for permanent and transitory shocks (where

[10]The negative coefficient is in line with the finding by Hunnes (2012) for Norwegian white-collar workers that demotions are associated with a decrease in real wages.

Table 3. Income Growth and Career Mobility

	Men			*Women*		
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Dependent variable: Change in real log labor income: $\ln(Y_t)-\ln(Y_{t-1})$						
Cross-firm move	0.011***			0.012***		
	(0.001)			(0.001)		
Upward move (PWF or PCF)		0.017***			0.013***	
		(0.003)			(0.004)	
Downward move (DWF or DCF)		–0.007*			–0.004	
		(0.004)			(0.005)	
Nonexecutive lateral move, within-firm (Stayer)			—			—
Nonexecutive lateral move, cross-firm (CF)			0.009***			0.012***
			(0.001)			(0.001)
Promotion, within-firm (PWF)			0.008***			0.009**
			(0.002)			(0.004)
Promotion, cross-firm (PCF)			0.064***			0.039***
			(0.012)			(0.010)
Executive no move (ExecStayer)			0.006***			0.005***
			(0.001)			(0.001)
Executive lateral move, cross-firm (ExecCF)			0.034***			0.019***
			(0.003)			(0.006)
Demotion, within-firm (DWF)			–0.006			–0.001
			(0.004)			(0.005)
Demotion, cross-firm (DCF)			–0.003			–0.008
			(0.008)			(0.015)
Age/10	–0.035***	–0.036***	–0.036***	–0.005	–0.006	–0.006
	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)	(0.005)
$\text{Age}^2/100$	0.002***	0.002***	0.002***	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)
Constant	0.118***	0.121***	0.121***	0.041***	0.043***	0.042***
	(0.007)	(0.007)	(0.007)	(0.011)	(0.011)	(0.011)
R^2	0.0065	0.0062	0.0072	0.0038	0.0031	0.0039
Number of observations	627,110	627,110	627,110	289,872	289,872	289,872
Number of unique individuals	57,010	57,010	57,010	26,352	26,352	26,352

Notes: OLS regression results for a set of models in which changes in real log labor income are regressed on different sets of employee-mobility dummies. All regressions include education, sector, and year dummies; they ignore transitory shocks. Clustered standard errors are reported in parentheses. OLS, ordinary least squares.
Significance levels: *** 1%; ** 5%; * 10%.

a permanent shock is understood as a random walk) to the earnings process. Accordingly, the residual in Equation (1) will contain both permanent and transitory components, and we have that

$$u_{it} = \upsilon_{it} + \Delta\varepsilon_{it}$$

where υ_{it} is an independent and identically distributed (iid) permanent income shock and ε_{it} is a transitory shock that follows a moving-average model of order q, MA(q), process. This implies that u_{it} will have nonzero autocorrelations up to order $q + 1$. Studies of individual earnings dynamics have typically found a low-order MA structure, suggesting that q should be around 2 (e.g., Abowd and Card 1989; Meghir and Pistaferri 2004).

Theoretically, the transitory shocks should be correlated with the mobility dummies and, specifically, with the promotion variables. To better understand this, consider the mean-regression model in Lazear (2004), in which productivity is determined, in part, by a transitory component. Although the expectation of this component is zero when taken over the population, among those employees who are promoted it should be positive because workers tend to exert higher effort prior to promotion or to being recruited by another firm, which will boost the transitory components of pay (e.g., bonuses and commissions). Hence, regression to the mean in the transitory component should reduce earnings growth somewhat in the year following many types of mobility, suggesting that the term $\Delta\varepsilon_{it}$ should be negatively correlated with the mobility and the promotion variables. Our estimation results show that larger coefficients are indeed obtained on the promotion variables once the covariance structure of the residual is properly accounted for.

In Appendix A, we show how the inclusion of lagged earnings growth can reduce the bias from omitted transitory shocks. Once we do this, the model becomes:

$$\Delta \ln\left(Y_{it}\right) = \alpha + \sum\nolimits_{s=1}^{S} \gamma_s \Delta \ln\left(Y_{i(t-s)}\right) + \sum\nolimits_{j=1}^{J} \mu_j M_{jit} + X'_{it}\beta + e_{it} \tag{2}$$

where the parameters γ_s reflect the correlation between the lagged earnings growth and transitory earnings shocks. The lag length S is chosen so that the e_{it} exhibit no significant serial correlation. In this sense, our specification is consistent with Abowd and Card (1989), Topel and Ward (1992), and Meghir and Pistaferri (2004), who modeled income as containing both permanent and transitory shocks.

The Econometric Treatment of Mobility

Our econometric treatment of mobility adds to important previous contributions in this literature (e.g., Topel and Ward 1992) by relaxing the strict exogeneity assumption and imposing, instead, the following moment conditions:

$$E\left[e_{it} M_{jis}\right] = 0 \text{ for } t \geq s \text{ and } \forall j$$

These conditions amount to assuming that mobility is predetermined because the residual in Equation (2) at time *t* is orthogonal to all mobility dated *t* and prior to *t*. This implies that the permanent income innovation embedded in e_{it} is allowed to affect mobility at $t + 1$ and beyond. As discussed in Arellano and Honoré (2001), our predeterminedness assumption restricts the serial correlation in e_{it}. This further emphasizes the importance of choosing the lag length *S* so that the residuals are serially uncorrelated to avoid issues of inconsistency. The residual e_{it} will contain all innovations to earnings growth that are "close to" being serially uncorrelated, where *close to* is precisely defined in Appendix A. In fact, for an *S* that is very large, e_{it} will be serially uncorrelated.

Having said that, even though our predeterminedness assumption is weaker than the strict exogeneity assumption and certainly at least as defensible as other identifying restrictions that have been employed elsewhere in the mobility literature, it is not innocuous. Causal interpretation is conditional on the moment conditions holding true. To the extent that earnings might affect contemporaneous mobility, for example, if job mobility is associated with wage offers as in the search literature, our results should be seen as descriptive. In that case, we might be tempted to invoke the even weaker assumption of endogeneity and apply lagged mobility as instruments for contemporaneous mobility in a generalized method of moments (GMM) framework. This, however, is not a viable alternative in our case because the lagged mobility variables constitute only weak instruments. The reason is that the amount of variation in these variables is limited because of the large number of flows we consider and the nature of our hierarchical placement measure. Thus, the predeterminedness assumptions appears to be necessary if we embed mobility measures into the framework of Meghir and Pistaferri (2004). Alternative instrumental variable (IV) strategies are not viable either. Although exogenous variations in the data could be exploited for some types of mobility (e.g., using job loss due to plant closure as an instrument), valid and strong instruments are not at hand for *all* eight types of mobility that we are considering.[11]

Does a Fixed Effect in Earnings Growth Exist?

A final issue related to our econometric strategy is whether we need to augment our earnings growth equation by allowing for a fixed effect α_i:

$$\Delta \ln\left(Y_{it}\right) = \alpha_i + \sum\nolimits_{s=1}^{S} \gamma_s \Delta \ln\left(Y_{i(t-s)}\right) + \sum\nolimits_{j=1}^{J} \mu_j M_{jit} + X'_{it}\beta + e_{it} \tag{3}$$

[11]Alternative ways of modeling promotion and wage dynamics are available. For example, Gibbons, Katz, Lemieux, and Parent (2005) developed an estimation strategy allowing for a time-invariant unobserved characteristic to drive comparative advantage and learning, and applied it to a study of mobility across sectors. Lluis (2005) applied a similar approach to within-firm wage and mobility dynamics. This approach differs from ours by assuming that the only stochastic driver of wage growth is the learning innovation, which is independent across time.

Table 4. Autocovariances of Income Growth

	Autocovariance				
Order	*0*	*1*	*2*	*3*	*4*
Men	0.02113***	−0.00627***	−0.00131***	−0.00036	−0.0000001
	(0.00210)	(0.00102)	(0.00052)	(0.00035)	(0.00010)
Women	0.01743***	−0.00437***	−0.00156	−0.00027	−0.00001
	(0.00290)	(0.00119)	(0.00094)	(0.00019)	(0.00004)

Notes: Autocovariances for men and women, used to discriminate among the statistical models. Bootstrap standard errors (100 replications) appear in parentheses.
Significance levels: *** 1%; ** 5%; * 10%.

In the presence of a fixed effect (i.e., if tests suggest that $\mathrm{Var}(\alpha_i) > 0$), we need to work with the earnings growth model in first differences:

$$\Delta\Delta \ln\left(Y_{it}\right) = \sum\nolimits_{s=1}^{S} \gamma_s \Delta\Delta \ln\left(Y_{i(t-s)}\right) + \sum\nolimits_{j=1}^{J} \mu_j \Delta M_{jit} + \Delta X'_{it}\beta + \Delta e_{it} \tag{4}$$

The double difference of log earnings then serves as the dependent variable, and the level of the mobility variables dated $t-1$ and earlier can be used as instruments for ΔM_{jit} in a GMM estimation (see Arellano and Bond 1991). For instance, Belzil and Bognanno (2008) used this procedure. If, however, $\mathrm{Var}(\alpha_i) = 0$, Equation (3) reduces to Equation (2), which we can directly estimate using OLS. Note that if $\mathrm{Var}(\alpha_i) > 0$, earnings growth will exhibit nonzero autocorrelations at arbitrarily long leads and lags.

Hence, the initial step in our analysis is to test whether $\mathrm{Var}(\alpha_i) > 0$. This will guide our choice of empirical model, selecting between a GMM procedure as in Arellano and Bond (1991) and OLS. Employing a procedure common in the earnings dynamics literature, our test is based on the autocorrelations of earnings growth (e.g., Abowd and Card 1989; Meghir and Pistaferri 2004). In the presence of a fixed effect in earnings growth, autocorrelations should be positive and significant at all leads and lags.

Table 4 reports the autocovariances along with their bootstrapped standard errors. These autocovariances are significantly negative up to order 2. This suggests that no fixed effect is present in earnings growth and that we can directly estimate Equation (3) under the assumption that the transitory earnings shocks are MA(1).[12]

A caveat to the results is that the test for the absence of a fixed effect in earnings growth in Table 4 can have low power. Baker (1997) illustrated this with an extract from the PSID that has approximately 500 individuals. But because we have more than 57,000 individuals in our male sample and more than 26,000 individuals in our female sample, we do not believe that this is an issue in our case. In addition, our balanced panel structure helps us avoid another potential problem that studies with panel data face: in

[12]Results are robust to assuming an MA process of higher order (available from the authors).

unbalanced panels, higher-order covariances are estimated with less data than lower-order ones, which can result in a failure to reject a false null of a zero autocovariance at high orders. To explore the robustness of our findings we, nevertheless, also estimate the specification in Equation (4) using GMM. Clearly, GMM and not OLS is needed here because the double differenced model contains both $e_{i(t-1)}$ and M_{jit} and our predeterminedness assumptions makes no guarantees that these will not be correlated.

Mobility and Earnings Growth: Estimation Results

Our main specification, based on Equation (2), yields the estimates reported in Table 5, columns (2) and (4), for men and women, respectively. Columns (1) and (3) allow comparisons with the previous estimates. Similar to previous contributions (e.g., da Silva and van der Klaauw 2011), we account for firm heterogeneity and time effects through sector and time dummies.

Consistent with our covariogram-based specification tests, the Cochrane–Orcutt test suggests that one lag of the dependent variable is sufficient to eliminate autocorrelation in the errors.[13] Lagged compensation growth has a negative effect on current compensation growth, a common finding in the earnings dynamics literature (e.g., Abowd and Card 1989; Topel and Ward 1992; Meghir and Pistaferri 2004). The negative serial correlation reflects the effects of transitory shocks. To the extent that high (low) earnings growth in the past period was driven by transitory productivity shocks and pay for performance, the regression will tend to move to the mean and lower (higher) earnings growth in the current period. In line with this explanation, Belzil and Bognanno (2008) attributed the negative serial correlation in their estimates for the overall earnings growth of U.S. executives to variable pay components.

Comparing our estimates with the biased specifications in Table 5, columns (1) and (3), upward mobility and cross-firm moves at the executive level have higher returns, whereas no change appears in the effect of cross-firm mobility at the nonexecutive level (CF) and the demotion coefficients remain insignificant. Both the male and female samples exhibit this pattern.

Our results reveal an asymmetry between the effect of a promotion and a demotion on wage growth: Both men and women gain more from moving up to an executive-level position than they lose when stepping down from such a position. The impact of demotions on wage growth has received little attention, except from Belzil and Bognanno (2008). Their study focused on reporting levels *within* the executive tier at 600 large U.S. firms from 1981 to 1988 and found that demotions have a stronger (negative) effect on compensation growth than promotions.

Exploiting the unique feature of our data that allows us to follow individuals across firm boundaries, we show that a great deal of heterogeneity exists in the returns to cross-firm moves. A cross-firm promotion leads to 3 to 6%

[13]Results are robust to including further lags, though (available from the authors).

Table 5. Income Growth and Career Mobility

	Men		Women	
	(1)	(2)	(3)	(4)
Labor income growth ($t-1$)		−0.309***		−0.304***
		(0.020)		(0.041)
Nonexecutive lateral move, within-firm (Stayer)	—	—	—	—
Nonexecutive lateral move, cross-firm (CF)	0.009***	0.009***	0.012***	0.012***
	(0.001)	(0.001)	(0.001)	(0.001)
Promotion, within-firm (PWF)	0.008***	0.012***	0.009**	0.015***
	(0.002)	(0.002)	(0.004)	(0.004)
Promotion, cross-firm (PCF)	0.064***	0.071***	0.039***	0.047***
	(0.012)	(0.012)	(0.010)	(0.010)
Executive no move (ExecStayer)	0.006***	0.009***	0.005***	0.007***
	(0.001)	(0.001)	(0.001)	(0.002)
Executive lateral move, cross-firm (ExecCF)	0.034***	0.039***	0.019***	0.023***
	(0.003)	(0.003)	(0.006)	(0.007)
Demotion, within-firm (DWF)	−0.006	−0.005	−0.001	−0.004
	(0.004)	(0.004)	(0.005)	(0.005)
Demotion, cross-firm (DCF)	−0.003	−0.002	−0.008	−0.012
	(0.008)	(0.007)	(0.015)	(0.015)
Age/10	−0.036***	−0.045***	−0.006	−0.007
	(0.003)	(0.004)	(0.005)	(0.006)
$\text{Age}^2/100$	0.002***	0.003***	0.000	0.000
	(0.000)	(0.000)	(0.001)	(0.001)
Constant	0.121***	0.152***	0.042***	0.061***
	(0.007)	(0.010)	(0.011)	(0.014)
R^2	0.0072	0.0975	0.0039	0.0809
Number of observations	627,110	570,100	289,872	263,520
Number of unique individuals	57,010	57,010	26,352	26,352
Cochran–Orcutt test (H_0: No autocorrelation in errors)	−0.308	−0.062	−0.303	−0.071
p value	< 0.001	0.152	<0.001	0.326

Notes: OLS regression results that account for transitory shocks, in which changes in real log labor income are regressed on different sets of employee mobility dummies. All regressions include education, sector, and year dummies. Clustered standard errors are reported in parentheses. OLS, ordinary least squares. Significance levels: *** 1%; ** 5%; * 10%.

faster growth than a within-firm promotion (with men gaining more than women do). A within-firm promotion, in turn, yields roughly the same return as a move across firm boundaries within the nonexecutive layer, both adding around 1% to earnings growth. At the executive level, a bigger gender difference appears; women gain around 2 percentage points from switching employers, whereas men gain around 4 percentage points. Finally, a cross-firm demotion lowers earnings growth to the level of a Stayer at the nonexecutive level.

Overall, we find sizable short-run gains from cross-firm moves, even after controlling for unobserved individual heterogeneity, in line with previous research on between-job earnings growth (e.g., Topel and Ward 1992; and using Danish data, Bagger et al. 2014). For example, the immediate growth premium associated with a lateral move across firms at the nonexecutive level is comparable to that from being promoted within the firm, which is in line with the finding by da Silva and van der Klaauw (2011), using Portuguese data, which showed that promotions and cross-firm transitions provide similar one-off gains. Our novel contribution is to show that the interaction effects with the hierarchical dimension account for a great deal of heterogeneity in the returns to cross-firm mobility; the gains from vertical moves in the hierarchy tend to be bigger if they are across firm boundaries than if they are within-firm.

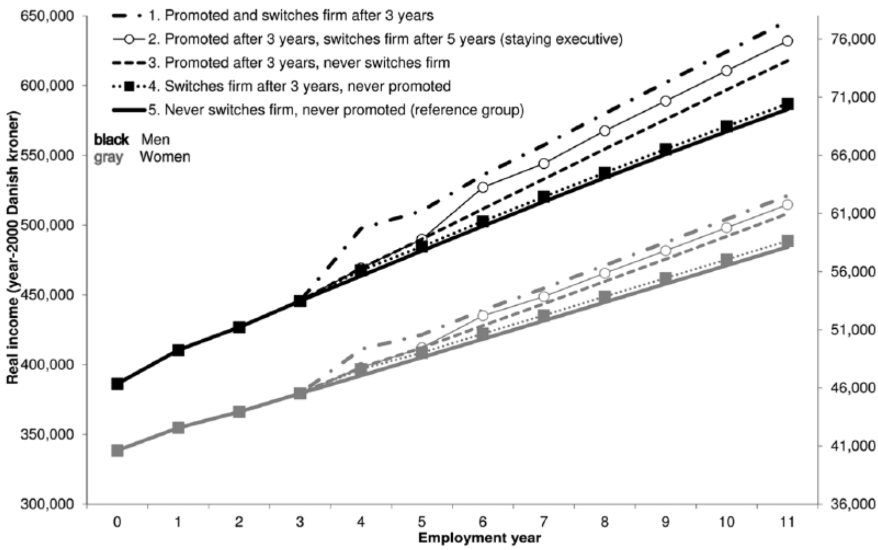
Implications for Earnings Growth Dynamics

What do our estimates imply for earnings growth after the various employment histories? The answer is not straightforward from the mobility coefficients in Table 5 because they paint only the short-run picture. To gauge the medium-run effects implied by our estimates, we compute the cumulative growth rates for the employment history scenarios from Table 5, columns (2) and (4). All the scenarios are based on the career of a university graduate with postgraduate education (17 years of education) who started employment at age 30. We compare a benchmark no-move scenario with the employment histories that involve a within-firm promotion or some type of cross-firm move (in our sample, relatively few workers switch employers repeatedly in a 10-year window). Figure 1 illustrates the resulting earnings paths and gives the quickest overview of the patterns that emerge (black lines refer to men and gray lines to women). Tables 6 and 7 provide a detailed analysis of log growth patterns for men and women with bootstrapped standard errors, respectively (see Appendix B for details on the bootstrapping procedure). We should keep in mind our definition of promotion as a move from the nonexecutive to the executive level. That is, we are not considering the total returns to any type of promotion over a career but, rather, the returns to this specific type of promotion.

The most striking feature that appears is that implied earnings outcomes after 10 years split neatly into the two categories: “never promoted” and “promoted.” Cross-firm mobility has a secondary effect only. The lowest placed black line in Figure 1 represents the reference group: a male employee who makes neither a vertical nor a cross-firm move (scenario 5). The second-lowest line represents an employee who switches employers after the third employment year but who is never promoted (scenario 4). Hence, the distance between the two bottom lines reflects the return to cross-firm mobility at the nonexecutive level. The top three lines represent career histories involving a promotion (scenarios 1 to 3). Comparing these lines, we can see that the gain from moving up to the executive level exceeds by far the gain from just switching employers. Within the group of promoted workers, we see that cross-firm mobility at the executive level provides sizable extra earnings growth (scenarios 1 and 2 compared to 3) but contributes less than the initial change in hierarchy levels.

The same pattern of relative growth rates also emerges from the separate estimation for the female sample (gray lines in Figure 1). A gender

Figure 1. Comparison of Real Income Growth of Men and Women



Notes: Evolution of real labor income implied by specifications in Table 5, columns (2) and (4), for a university graduate (Education = 17) starting his or her career at age 30.

difference does exist, however, in the average starting salary and absolute growth rates. The latter can be seen more clearly in Figure 2, which plots the earnings indices for men and women, which reflect how earnings grow relative to the level at the start of the career. For our purposes, the important message is that the two separate samples yield a consistent picture: that cross-firm mobility offers more modest gains than vertical mobility. Given the differences between men and women (e.g., in labor force participation and fertility considerations), the robustness of our findings across samples is quite remarkable.

Tables 6 and 7 quantify the effects and tell us that the differential growth rates are indeed statistically significant. The top part of the tables shows how much the real income of a 30-year-old university graduate is predicted to grow over a 5-year and 10-year horizon. For example, a male employee who was promoted after three years but never switched employers (scenario 3) has a predicted real income growth of around 63% over 10 years ($\exp(0.487) \approx 1.63$), whereas someone who switched employers after 3 years but was never promoted (scenario 4) sees real income growth of around 55% ($\exp(0.435) \approx 1.55$). The earnings growth is lower than the estimates for the United States that control for individual fixed effects. For example, Schönberg (2007) reported a 10-year growth rate of around 80% for university graduates, and Topel and Ward (1992) found that earnings roughly double. Note, however, that these figures are hard to compare because the U.S. studies look at the early stages of a career and include individuals with a weaker labor-force attachment than in our sample.

The bottom part of Tables 6 and 7 compares the real earnings growth across career histories. Continuing with our example, consider the comparisons of the scenarios that involve a promotion and those that do not. We see that the

Table 6. Income Growth Dynamics, Men

A. Scenarios and growth

Career scenarios	*Cumulative log income growth after* Year 5	Year 10
1. Promoted and switched firm after 3 years	0.322	0.533
2. Promoted after 3 years, switched firm after 5 years (staying executive)	0.305	0.510
3. Promoted after 3 years, never switched firm	0.275	0.487
4. Switched firm after 3 years, never promoted	0.257	0.435
5. Never switched firm, never promoted (reference group)	0.250	0.428

B. Comparisons

	Year 5 difference in		*Year 10 difference in*		
	Cumulative log income growth	*Current log income growth*	*Cumulative log income growth*	*Current log income growth*	
1 and 5	0.071***	0.013***	0.103***	0.007***	
	(0.007)	(0.001)	(0.009)	(0.001)	Promoted compared to never promoted
2 and 5	0.054***	0.037***	0.080***	0.007***	
	(0.003)	(0.003)	(0.005)	(0.001)	
3 and 5	0.024***	0.007***	0.058***	0.007***	
	(0.002)	(0.001)	(0.005)	(0.001)	
1 and 4	0.064***	0.012***	0.097***	0.007***	
	(0.008)	(0.001)	(0.009)	(0.001)	
2 and 4	0.047***	0.036***	0.074***	0.007***	
	(0.003)	(0.003)	(0.005)	(0.001)	
3 and 4	0.018***	0.006***	0.051***	0.007***	
	(0.002)	(0.001)	(0.005)	(0.001)	
1 and 3	0.047***	0.006***	0.045***	0.000***	Within promoted group
	(0.007)	(0.001)	(0.007)	(0.000)	
1 and 2	0.017*	−0.024***	0.023***	0.000***	
	(0.009)	(0.003)	(0.008)	(0.000)	
2 and 3	0.029***	0.029***	0.022***	0.000***	
	(0.003)	(0.003)	(0.002)	(0.000)	
4 and 5	0.007***	0.001***	0.007***	0.000	Within never promoted group
	(0.001)	(0.000)	(0.001)	(0.000)	

Notes: Simulated income growth for men with different mobility patterns. Predictions are based on Table 5, column (2), and assume a starting age of 30 and Education = 17. Bootstrap standard errors are reported in parentheses (100 replications).
Significance levels: *** 1%; ** 5%; * 10%.

10-year income under scenario 3 is about 6% higher than it would have been under scenario 5 ($\exp(0.058) \approx 1.06$). A comparison of the 10-year log growth rates and their components suggests that 12 to 18% of the total growth can be attributed to promotions and only 2 to 9% can be attributed to gains from cross-firm mobility.[14]

[14]Of the total earnings growth in scenario 1 in the male sample, 0.533, a comparison with scenario 4 shows that 0.098 log points (18%) can be attributed to the promotion. Comparing scenarios 1 and 3, we can see that 0.046 log points (9%) can be attributed to the cross-firm move. Similarly, comparisons of scenarios 2 and 3 and of scenarios 4 and 5 attribute 2 to 5% of the 10-year earnings growth to cross-firm mobility. And a comparison of scenarios 3 and 5 attributes 12% to promotions. Figures for the female sample are obtained in similar fashion.

Table 7. Income Growth Dynamics, Women

A. Scenarios and growth

Career scenarios	Cumulative log income growth after Year 5	Cumulative log income growth after Year 10
1. Promoted and switched firm after 3 years	0.256	0.437
2. Promoted after 3 years, switched firm after 5 years (staying executive)	0.247	0.425
3. Promoted after 3 years, never switched firm	0.231	0.412
4. Switched firm after 3 years, never promoted	0.218	0.374
5. Never switched firm, never promoted (reference group)	0.208	0.365

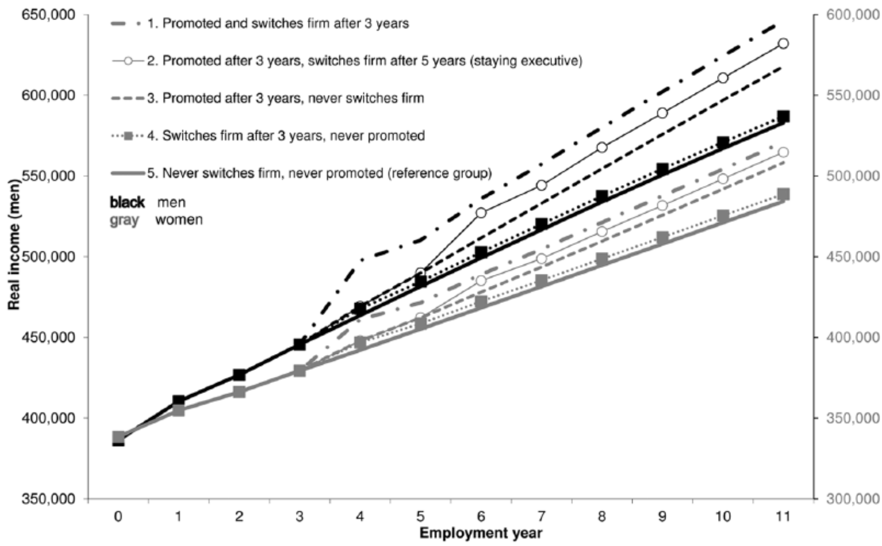
B. Comparisons

	Year 5 difference in		*Year 10 difference in*		
	Cumulative log income growth	*Current log income growth*	*Cumulative log income growth*	*Current log income growth*	
1 and 5	0.048***	0.009***	0.072***	0.005***	Promoted compared to never promoted
	(0.009)	(0.002)	(0.012)	(0.001)	
2 and 5	0.039***	0.022***	0.060***	0.005***	
	(0.007)	(0.006)	(0.009)	(0.001)	
3 and 5	0.022***	0.006***	0.047***	0.005***	
	(0.004)	(0.001)	(0.009)	(0.001)	
1 and 4	0.039***	0.008***	0.063***	0.005***	
	(0.009)	(0.002)	(0.012)	(0.001)	
2 and 4	0.030***	0.021***	0.051***	0.005***	
	(0.007)	(0.006)	(0.009)	(0.001)	
3 and 4	0.013***	0.003***	0.038***	0.005***	
	(0.004)	(0.001)	(0.009)	(0.001)	
1 and 3	0.026***	0.003***	0.025***	0.000	Within promoted group
	(0.010)	(0.001)	(0.009)	(0.000)	
1 and 2	0.009	–0.014***	0.012	0.000	
	(0.011)	(0.006)	(0.010)	(0.000)	
2 and 3	0.017***	0.017***	0.013***	0.000	
	(0.006)	(0.006)	(0.005)	(0.000)	
4 and 5	0.009***	0.001***	0.009***	0.000	Within never promoted group
	(0.001)	(0.000)	(0.001)	(0.000)	

Notes: Simulated income growth for women with different mobility patterns. Predictions are based on Table 5, column (2), and assume a starting age of 30 and Education = 17. Bootstrap standard errors are reported in parentheses (100 replications).
Significance levels: *** 1%; ** 5%; * 10%.

The overall picture is that all the promotion to no promotion scenarios comparisons yield greater differences in the 10-year cumulative growth rates than the within-group comparisons (Tables 6 and 7). At the 5-year horizon, the transitory effects of the mobility after the third employment year still lead to different short-run earnings growth rates. At the 10-year horizon, promoted workers (scenarios 1, 2, and 3) now are all on a significantly steeper growth path than never promoted workers (scenarios 4 and 5); earnings grow 0.7% faster. This is in line with learning models such as

Figure 2. Comparison of Relative Income Growth of Men and Women



Notes: Evolution of real labor income implied by specifications in Table 5, columns (2) and (4), for a university graduate (Education = 17) starting his or her career at age 30.

Gibbons and Waldman (1999b, 2006), in which the assignment to a higher-level job entails a steeper earnings growth path.

Robustness Checks and Further Analysis

In this section we show that our findings are robust to estimating on subsamples with different education levels, allowing for individual-level trends in earnings growth, and that they are not sensitive to the firm-size restriction used to obtain our core sample. We also investigate the sensitivity of our results to mobility across industry boundaries and to mobility to firms of a significantly different size (larger or smaller). Finally, we investigate the variation in earnings outcomes that are attributable to particular mobility patterns.

Estimations on Subsamples with Different Education Levels

Our main estimation results control for differences in education using dummies that distinguish four categories. The group with nine years of education have completed only the compulsory schooling (omitted category, 17% of the sample). Those with 12 years of schooling have a high school diploma (56% of the sample). The group with 15 years of schooling includes those with a bachelor's degree or who have completed an apprenticeship or some other form of post–secondary school professional training (19% of the sample). The final category, with 17 years or more of schooling, includes those with a postgraduate university education, having completed a master's degree or doctorate (8% of the sample). Although our main estimation allows for different growth rates across education categories, it restricts the returns to mobility to being the same for all education levels. Table 8 shows

Table 8. Education Subsamples, Men

	Full sample	*Education = 9*	*Education = 12*	*Education = 15*	*Education = 17*
Dependent variable: Change in real log labor income: $\ln(Y_t)-\ln(Y_{t-1})$					
Labor income growth $(t-1)$	−0.309***	−0.352***	−0.270***	−0.344***	−0.306***
	(0.020)	(0.041)	(0.032)	(0.032)	(0.054)
Nonexecutive lateral move, within-firm (Stayer)	—	—	—	—	—
Nonexecutive lateral move, cross-firm (CF)	0.009***	0.002	0.005	0.019***	0.021***
	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)
Promotion, within-firm (PWF)	0.012***	−0.002	0.011***	0.018***	0.016***
	(0.002)	(0.005)	(0.003)	(0.003)	(0.006)
Promotion, cross-firm (PCF)	0.071***	0.029***	0.072***	0.068***	0.092***
	(0.012)	(0.011)	(0.025)	(0.011)	(0.036)
Executive no move (ExecStayer)	0.009***	0.006***	0.008***	0.010***	0.016***
	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)
Executive lateral move, cross-firm (ExecCF)	0.039***	0.033***	0.026***	0.050***	0.046***
	(0.003)	(0.009)	(0.005)	(0.006)	(0.007)
Demotion, within-firm (DWF)	−0.005	−0.005	−0.006	−0.003	−0.016
	(0.004)	(0.007)	(0.006)	(0.004)	(0.001)
Demotion, cross-firm (DCF)	−0.002	−0.038	−0.015	−0.022**	−0.001
	(0.007)	(0.028)	(0.010)	(0.011)	(0.025)
Age/10	−0.045***	0.023***	−0.034***	−0.058***	−0.119***
	(0.004)	(0.011)	(0.005)	(0.011)	(0.022)
$\text{Age}^2/100$	0.003***	−0.004***	0.002***	0.004***	0.009***
	(0.000)	(0.001)	(0.001)	(0.001)	(0.002)
Education = 9	-				
Education = 12	0.003***				
	(0.000)				
Education = 15	0.009***				
	(0.001)				
Education = 17	0.016***				
	(0.001)				
Constant	0.152***	−0.018	0.118***	0.219***	0.397***
	(0.010)	(0.023)	(0.012)	(0.023)	(0.049)
R^2	0.0974	0.1225	0.0788	0.1172	0.0925
Number of observations[a]	570,100	95,926	317,770	112,339	44,065
Cochran–Orcutt test (H_0: No autocorrelation in errors)	−0.062	−0.072	−0.042	−0.084	−0.078
p value	0.152	0.518	0.461	0.333	0.521

Notes: Education level measured in years of education. OLS regression results that account for transitory shocks for each education group. All regressions include sector and year dummies. Clustered standard errors are reported in parentheses. OLS, ordinary least squares.
[a]In the subsamples, a few individuals increased their education level. For them, observations are split between the relevant education categories according to the education level in the respective year.
Significance levels: *** 1%; ** 5%; * 10%.

the estimation results for men when this assumption is relaxed and the subsamples are analyzed separately.

Overall, the results for men in Table 8 show that all point-estimates on the mobility dummies tend to increase as the education level increases, and Figure 3 illustrates this pattern very clearly. Employees with a postgraduate university degree (Figure 3d) have higher returns to mobility than those with a college degree or post–secondary school professional training (Figure 3c), who in turn gain more than those with a high school diploma (Figure 3b) or less education (Figure 3a). But in all cases, we again observe a divide between the promoted and never promoted categories, as in our main estimates.

A closer look at Table 8 indicates that a move to a new firm at the nonexecutive level and a within-firm promotion offer more or less equal short-term gains for employees with at least 15 years of education. Nevertheless, both cross-firm promotions and cross-firm mobility at the executive level remain the most lucrative types of mobility. For employees with a high school diploma, a promotion boosts growth at almost twice the rate associated with cross-firm mobility at the nonexecutive level. For the highly educated employees, the returns from cross-firm promotions and mobility at the executive level are highest. For the lowest education group, the coefficients are less precisely estimated because some flows have few observations. For women, the patterns are similar but less pronounced.[15]

Overall, the qualitative results are in line with those for the full sample. Splitting the sample by education, however, does reveal additional details. Our main specification captures differences in overall earnings growth across the education groups through education dummies. When we estimate subsample by subsample, the differences in growth rates across the education groups show up in the regression constants. The different slopes of earnings growth in Figure 3 reflect this. The growth rates clearly increase with the level of education. For instance, in the 10-year earnings growth for the base scenario (scenario 5), employees with no high school diploma achieve a pay progression of around 10%, whereas university graduates more than double their earnings.

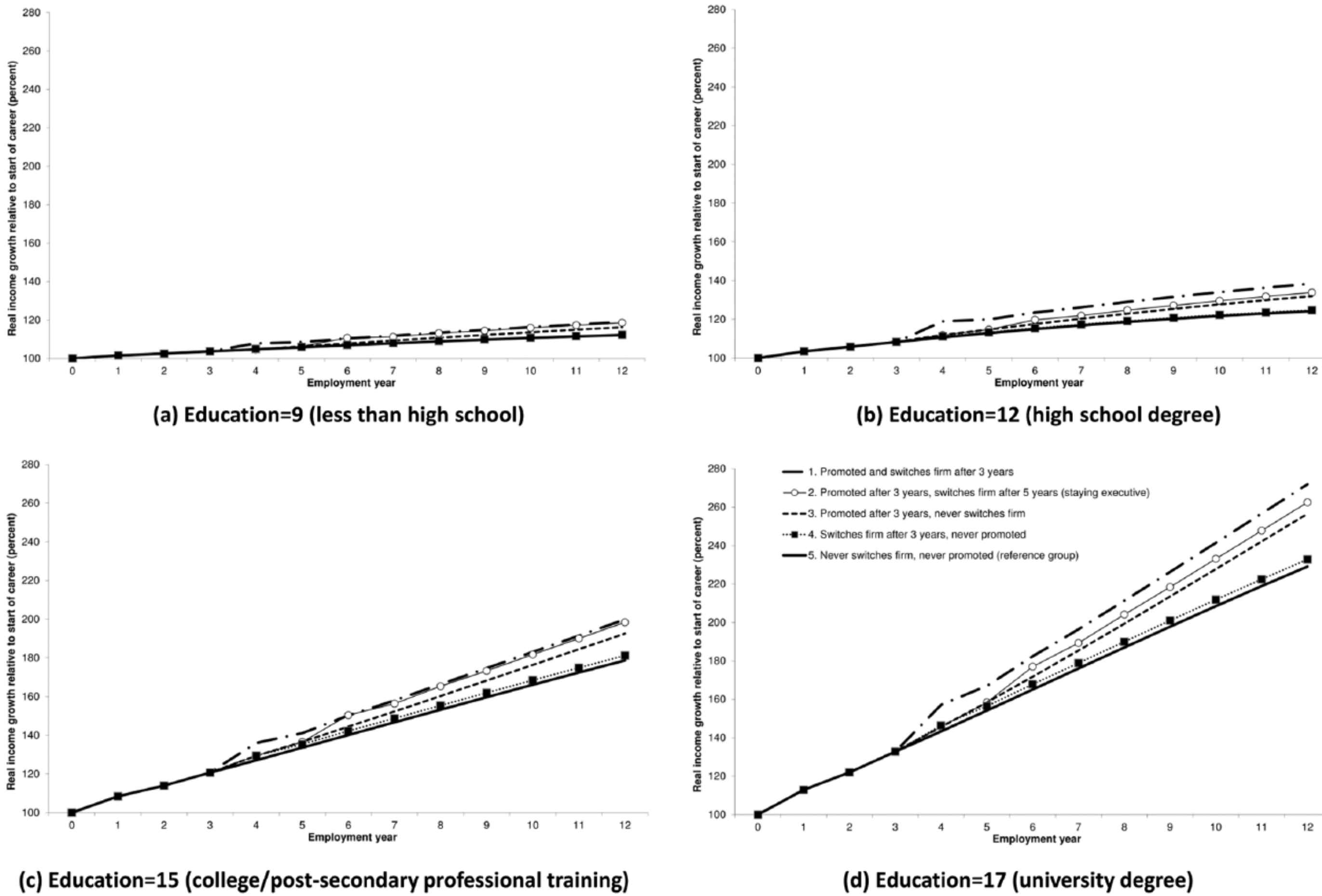
Evaluating the Importance of the Firm-Size Restriction

In our main analysis, we focus on a sample of core employees who work continuously in firms with at least 25 employees. To explore whether the gains from mobility are sensitive to this size restriction, we re-estimate our model using different criteria for inclusion in the sample. Table 9 presents the results for men. Column (1) restates the original results with a minimum firm size of 25, and columns (2) and (3) use the size restriction 50 and 100, respectively. Even though the sample size is reduced by up to 33% for men, the estimates are remarkably similar across samples. The results for women, in which the sample size is reduced by 25%, show the same patterns.[16] From this we conclude that our results are not driven by the firm-size criterion.

[15]Results are available from the authors.

[16]Results are available from the authors.

Figure 3. Relative Growth in Real Income for Men, by Education



Notes: Evolution of real labor income implied by Table 8 for a man starting his career at age 30.

Table 9. Robustness: Firm Size, Men

	Firm size		
	(1)	*(2)*	*(3)*
	≥ 25 (main sample)	*≥ 50*	*≥ 100*
Dependent variable: Change in real log labor income: $\ln(Y_t) - \ln(Y_{t-1})$			
Labor income growth $(t-1)$	−0.309***	−0.303***	−0.304***
	(0.020)	(0.024)	(0.028)
Nonexecutive lateral move, within-firm (Stayer)	—	—	—
Nonexecutive lateral move, cross-firm (CF)	0.009***	0.008***	0.006***
	(0.001)	(0.001)	(0.001)
Promotion, within-firm (PWF)	0.012***	0.015***	0.016***
	(0.002)	(0.002)	(0.002)
Promotion, cross-firm (PCF)	0.071***	0.069***	0.071***
	(0.012)	(0.014)	(0.019)
Executive no move (ExecStayer)	0.009***	0.010***	0.010***
	(0.001)	(0.001)	(0.001)
Executive lateral move, cross-firm (ExecCF)	0.039***	0.041***	0.040***
	(0.003)	(0.003)	(0.004)
Demotion, within-firm (DWF)	−0.005	−0.001	0.005
	(0.004)	(0.004)	(0.004)
Demotion, cross-firm (DCF)	−0.002	0.001	0.004
	(0.007)	(0.008)	(0.010)
Age/10	−0.045***	−0.046***	−0.041***
	(0.004)	(0.005)	(0.005)
$\text{Age}^2/100$	0.003***	0.003***	0.002***
	(0.000)	(0.001)	(0.001)
Constant	—	—	—
R^2	0.0974	0.0930	0.0967
Numbers of observations	570,100	467,620	377,620
Number of unique individuals	57,010	46,762	37,762
Cochran–Orcutt test (H_0: No autocorrelation in errors)	−0.062	−0.064	−0.055
p value	0.152	0.199	0.340

Notes: OLS regression results that account for transitory shocks for various firm-size restrictions. All regressions include sector and year dummies. Clustered standard errors are reported in parentheses. OLS, ordinary least squares.
Significance levels: *** 1%; ** 5%; * 10%.

Table 10 looks at the effects of moving to a new firm of a different size, interacting the mobility dummies with a dummy for moving to a firm that is either twice or half the size. Overall, a premium seems to exist for cross-firm moves to firms of a substantially different size, increasing real earnings growth by around 1 to 2 percentage points. Disaggregating by mobility type, we see that lateral cross-firm moves to smaller firms, in particular, carry a premium. For vertical movements, we see no such firm-size patterns (except for women who gain when being promoted when moving to a smaller firm). Small firms often occupy niches that require specialized expertise. The results indicate that as a result of the matching process small firms may offer a premium to employees with complementary skills and experience. Likewise, if we interact the mobility dummies with a dummy for moving to a new

Table 10. Income Growth and the Effects of Moving to a New Firm of a Different Size

	Men		*Women*	
	(1)	*(2)*	*(3)*	*(4)*
Dependent variable: Change in real log labor income: $ln(Y_t) - ln(Y_{t-1})$				
Labor income growth ($t-1$)	−0.309***	−0.309***	−0.304***	−0.304***
	(0.020)	(0.020)	(0.041)	(0.041)
Move to a firm at least twice the size of the previous one between $t-1$ and t (Bigger)	0.009***		0.021***	
	(0.002)		(0.002)	
Move to a firm less than half the size of the previous one between $t-1$ and t (Smaller)	0.022***		0.023***	
	(0.002)		(0.002)	
Other cross-firm move between $t-1$ and t	0.003***		0.002	
	(0.001)		(0.001)	
Nonexecutive lateral move, within-firm (stayer)		—		—
Nonexecutive lateral move, cross-firm (CF)		0.002*		0.001*
		(0.001)		(0.001)
Bigger*CF		0.009***		0.022***
		(0.002)		(0.003)
Smaller*CF		0.021***		0.023***
		(0.002)		(0.002)
Executive no move (ExecStayer)		0.009***		0.007***
		(0.001)		(0.002)
Executive lateral move, cross-firm (ExecCF)		0.031***		0.024***
		(0.004)		(0.009)
Bigger*ExecCF		0.003		−0.022
		(0.008)		(0.014)
Smaller*ExecCF		0.027***		0.023***
		(0.009)		(0.018)
Promotion, within-firm (PWF)		0.012***		0.015***
		(0.002)		(0.004)
Promotion, cross-firm (PCF)		0.056***		0.020***
		(0.017)		(0.015)
Bigger*PCF		0.027		0.019
		(0.037)		(0.021)
Smaller*PCF		0.020		0.077***
		(0.020)		(0.025)
Demotion, within-firm (DWF)		−0.005		−0.005
		(0.004)		(0.005)
Demotion, cross-firm (DCF)		−0.008		0.010
		(0.012)		(0.013)
Bigger*DFC		0.010		−0.023
		(0.018)		(0.034)
Smaller*DFC		0.011		−0.054
		(0.018)		(0.039)
R^2	0.097	0.098	0.081	0.082
Number of observations	570,100	570,100	263,520	263,520

Notes: OLS regression results that account for transitory shocks and capture the effects of moving to a firm of a different size. All regressions control for lagged income growth; the individual's education, age, and age squared; and year and sector dummies. Clustered standard errors are reported in parentheses. OLS, ordinary least squares.
Significance levels: *** 1%; ** 5%; * 10%.

industry, we see a premium of about 2 percentage points (1 percentage point for women) extra earnings growth for moving to a new industry.[17]

Distinguishing Particular Patterns of Mobility

About 15% of the male sample experienced some sort of vertical movement during the 12-year sample window of their career. Of these, about half experienced just one vertical transition (about 5% of the sample experienced just a promotion and about 2% experienced just a demotion), whereas the other half experienced nonstandard career trajectories. The most common among these was the move from the nonexecutive to the executive level and back again (accounting for about 3.5% of the sample). Looking at cross-firm mobility, about 36% of men make no move, 27% make one move, and 24% make two moves. More than two cross-firm moves are relatively uncommon.

Table 11 reveals that having a career trajectory that ends at the executive level is important and the path by which the worker gets there matters less. Whether a worker is promoted just once or has been demoted intermittently and then promoted again, the career trajectories ending at the executive level are all associated with about 14 to 17% earnings growth ($\exp(0.128) \approx 1.14$ and $\exp(0.154) \approx 1.17$). Further, the earnings growth increases with the number of cross-firm moves, which is in line with the hypothesis that ambitious employees seek out better opportunities and move to get them.[18] This replicates the finding by Smeets (2006) that a worker's wage is positively correlated with the number of past cross-firm moves. She found a similar correlation pattern in the American National Longitudinal Study of Youth. Again, using data similar to ours, Bagger (2006) found that the starting wages in the new job increase with additional job-to-job transitions but that this effect is small relative to the gain obtained on the first job-to-job transition. Earlier evidence for the United States has suggested that persistent movers witness lower wage growth (Mincer and Jovanovic 1981).

Allowing for a Fixed Effect in Earnings Growth

Our choice of estimation procedure was guided by the fact that the autocorrelation in earnings growth dies off quickly and becomes insignificant after a few lags, which is at odds with a fixed effect in earnings growth. Nevertheless, as a robustness check, we allow for unobserved, persistent individual heterogeneity in earnings growth. Table 12 reports the corresponding GMM estimates. The consistency of the Arellano–Bond estimator relies on the residuals (in first differences) to be uncorrelated with lags on the right-hand-side regressors as well as restrictions on serial correlation in the residuals. For the reported specifications, this requires adding two lags of the dependent variable as regressors, as the test for autocorrelation developed by Arellano and Bond (1991) shows.

[17]Results are available from the authors.

[18]The probability of promotion also increases with cross-firm mobility during the previous eight years (see also Frederiksen and Kato 2011).

Table 11. Income Growth and the Effects of Particular Mobility Patterns

	Men	*Women*
	(1)	*(2)*
Dependent variable: Change in real log labor income: $\ln(Y_{2005}) - \ln(Y_{1994})$		
No mobility	—	—
Vertical mobility ending at executive level		
Executive → Nonexecutive → Executive	0.128***	0.113***
	(0.013)	(0.041)
Nonexecutive → Executive	0.138***	0.134***
	(0.006)	(0.012)
Nonexecutive → Executive → Nonexecutive → Executive	0.154***	0.132***
	(0.013)	(0.029)
Vertical mobility ending at nonexecutive level		
Executive → Nonexecutive	0.002	0.007
	(0.012)	(0.018)
Executive → Nonexecutive → Executive → Nonexecutive	0.030	0.060
	(0.020)	(0.066)
Nonexecutive → Executive → Nonexecutive	0.042***	0.037*
	(0.007)	(0.021)
More complex vertical mobility patterns	0.063***	0.007
	(0.016)	(0.039)
Cross-firm mobility		
One cross-firm move	–0.002	–0.001
	(0.003)	(0.004)
Two cross-firm moves	0.005*	0.006
	(0.003)	(0.004)
Three cross-firm moves	0.010**	0.013*
	(0.004)	(0.007)
Four cross-firm moves	0.013**	0.023**
	(0.006)	(0.009)
More than four cross-firm moves	0.021*	–0.011
	(0.013)	(0.080)
R^2	0.134	0.050
Number of observations	57,010	26,352

Notes: OLS regression results that account for transitory shocks and capture the effects of particular mobility patterns. All regressions control for the individual's education, age, and age squared in 1994. Clustered standard errors are reported in parentheses. OLS, ordinary least squares.
Significance levels: *** 1%; ** 5%; * 10%.

Figure 4 illustrates the GMM results. It plots the evolution of the real labor income implied by the estimates for a university graduate starting his career at age 30. Figure 4 should be compared with Figure 1, which is based on our preferred specification. Again, the evolution of earnings depends mostly on whether a person manages to move up in the hierarchy and, to a lesser extent only, on cross-firm mobility. With the GMM estimation, the divide between the two categories, never promoted and promoted, even becomes larger.

Conclusion

Two well-developed yet largely separate literatures exist, on cross-firm mobility and on mobility within the hierarchy of organizations. The descriptive

Table 12. Labor Income Growth and Career Mobility

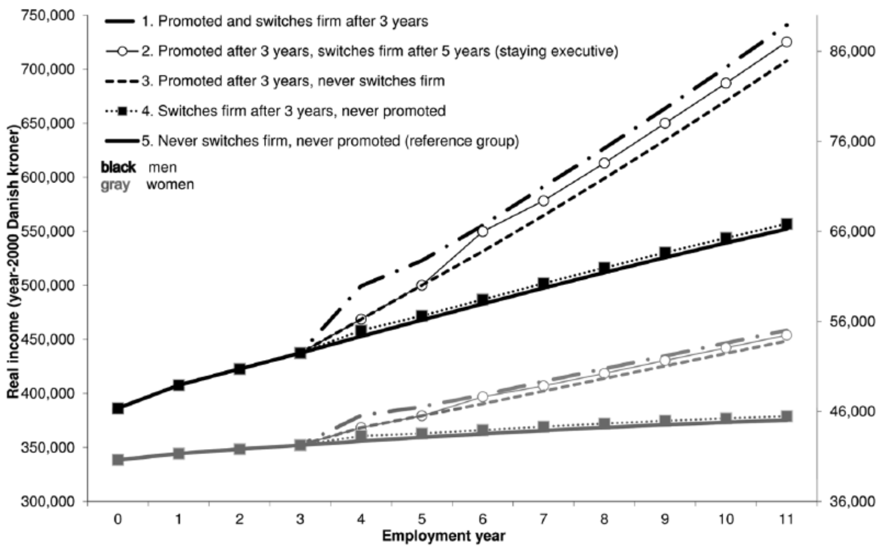
	Men		*Women*	
	OLS	*GMM*	*OLS*	*GMM*
Dependent variable: Change in real log labor income: $\ln(Y_t) - \ln(Y_{t-1})$				
Labor income growth $(t-1)$	−0.309***	−0.297***	−0.304***	−0.268***
	(0.020)	(0.020)	(0.041)	(0.033)
Labor income growth $(t-2)$		−0.097***		−0.111***
		(0.011)		(0.021)
Nonexecutive lateral move, within-firm (Stayer)	—	—	—	—
Nonexecutive lateral move, cross-firm (CF)	0.009***	0.012***	0.012***	0.013***
	(0.001)	(0.001)	(0.001)	(0.002)
Promotion, within-firm (PWF)	0.012***	0.035***	0.015***	0.035***
	(0.002)	(0.004)	(0.004)	(0.011)
Promotion, cross-firm (PCF)	0.071***	0.098***	0.047***	0.064***
	(0.012)	(0.009)	(0.010)	(0.016)
Executive no move (ExecStayer)	0.009***	0.042***	0.007***	0.029***
	(0.001)	(0.005)	(0.002)	(0.012)
Executive lateral move, cross-firm (ExecCF)	0.039***	0.076***	0.023***	0.045***
	(0.003)	(0.006)	(0.007)	(0.014)
Demotion, within-firm (DWF)	−0.005	0.010***	−0.004	−0.006
	(0.004)	(0.003)	(0.005)	(0.007)
Demotion, cross-firm (DCF)	−0.002	0.028***	−0.012	−0.003
	(0.007)	(0.008)	(0.015)	(0.030)
Age/10	−0.045***		−0.007	
	(0.004)		(0.006)	
$\text{Age}^2/100$	0.003***	−0.003***	0.000	−0.001***
	(0.000)	(0.000)	(0.001)	(0.000)
Constant	0.152***	0.068***	0.061***	0.030
	(0.010)	(0.017)	(0.014)	(0.051)
Number of observations	570,100	513,090	263,520	237,168
Number of unique individuals	57,010	57,010	26,352	26,352
Number of instruments	—	380	—	380
m_1		−6.811		−3.557
p value		< 0.001		<0.001
m_2		−0.200		−0.025
p value		0.842		0.980

Notes: Benchmarks of the OLS regression results that account for transitory shocks and GMM regression results. OLS regressions include education, sector, and year dummies. Clustered standard errors are reported in parentheses. GMM, generalized method of moments; OLS, ordinary least squares.
Significance levels: *** 1%; ** 5%; * 10%.

evidence on the effects of within- and cross-firm mobility on earnings growth that we have provided, based on Danish matched employer–employee panel data, informs research that considers both job search and hierarchical placement as drivers of wage growth.

Our estimates reveal sizable short-run gains for cross-firm mobility at the nonexecutive level. Yet the bulk of the longer-term earnings growth that we observe appears to be driven by promotions (either within or across firms) or to be a consequence of cross-firm mobility at the executive level. We have also established the existence of a substantial heterogeneity in pay progression

Figure 4. Illustration of Real Income Growth for Men and Women, GMM Estimates



Notes: Evolution of real labor income implied by the GMM estimates in Table 12 for a university graduate (Education = 17) starting his or her career at age 30.

between executives and nonexecutives, which is consistent with models of job assignment in which a promoted employee is placed in a position with a steeper earnings growth trajectory (e.g., Bernhardt 1995; Gibbons and Waldman 1999b, 2006). Overall, our results show that researchers need to consider both cross-firm mobility and hierarchical transitions, and to pay close attention to the interaction effects between these types of flows, to achieve a full understanding of the way mobility influences earnings progression.

We used the framework of Meghir and Pistaferri (2004) to analyze job mobility and wage growth. A caveat of this approach is that it requires a predeterminedness assumption for identification, as discussed in The Econometric Treatment of Mobility subsection. Our attempts to invoke the weaker endogeneity assumption, in which mobility is allowed to be freely correlated with wage offers (and, hence, wage growth), resulted in weak instruments. In our view, overcoming this hurdle is the main challenge that researchers face. Once comprehensive models that incorporate both search and hierarchical moves appear in the literature, the structural estimation of such models may be the way to proceed. Thus, a promising avenue for future work is to develop an integrated theoretical framework that explicitly addresses the underlying determinants both of hierarchical transitions and of employee turnover, which then can be taken back to the data.[19]

[19]We are aware of ongoing work by Veramendi (2011). Using data similar to ours, he showed that the workhorse job-ladder search model has problems explaining worker mobility, and he used this as motivation for a new model, in which workers accumulate skills from learning by doing, with decreasing returns in a given job. To continue learning new skills, workers must either be promoted within their firm or search for outside opportunities. Together with Rune Vejlin, Veramendi is currently refining this model to structurally estimate it using Danish data similar to ours, in a project entitled "Search, Human Capital, and Career Development" (personal communication, June 3, 2014).

Appendix A

Econometric Details

Define $y_{it} = \ln Y_{it}$, and suppose that earnings growth is equal to the sum of the permanent and transitory innovations:

$$\Delta y_{it} = \upsilon_{it} + \Delta\varepsilon_{it} \equiv \zeta_{it}$$

where v_{it} is iid white noise and ε_{it} is a stationary MA(q) process. Note that ζ_{it} is an MA($q+1$) process and so can be written as:

$$\zeta_{it} = (1 - \theta_1 L^1 - \ldots - \theta_{q+1} L^{q+1})\tilde{e}_{it}$$

where $\tilde{e}_{it}$ is iid white noise. If we can invert the lag polynomial, then we obtain:

$$\tilde{e}_{it} = (1 - \theta_1 L^1 - \ldots - \theta_{q+1} L^{q+1})^{-1}\zeta_{it} \equiv \sum_{s=0}^{\infty} \tilde{\gamma}_s \Delta y_{i(t-s)}$$

where $\tilde{\gamma}_0 = 1$. This then implies that:

$$\Delta y_{it} = \sum_{s=1}^{\infty} \gamma_s \Delta y_{i(t-s)} + \tilde{e}_{it}$$

where $\gamma_s = -\tilde{\gamma}_s$. But because the process $\Delta y_{i,t}$ is ergodic, the coefficients γ_s will tend to zero for large s, and thus, very little is sacrificed by writing:

$$\Delta y_{it} = \sum_{s=1}^{S} \gamma_s \Delta y_{i(t-s)} + e_{it}$$

where $e_{i,t} = \sum_{s=S+1}^{\infty} \gamma_s \ \Delta y_{i(t-s)} + \tilde{e}_{i,t}$, which is essentially what we have written in the main part of the article. The only issue that remains is to show that the lag polynomial is invertible. We can show that it is indeed invertible in the special case where ε_{it} is iid white noise. This implies that ζ_{it} is MA(1), and so

$$\zeta_{it} = (1 - \theta L)\tilde{e}_{it}$$

We then obtain:

$$E\left(\Delta y_{it}^2\right) = \sigma_{\tilde{e}}^2\left(1 + \theta^2\right) = \sigma_{\upsilon}^2 + 2\sigma_{\varepsilon}^2 \tag{5}$$

$$E\left(\Delta y_{it} \Delta y_{i(t-1)}\right) = -\theta \tilde{e}_{it} = -\sigma_{\varepsilon}^2 \tag{6}$$

Equation (6) implies that $\theta = \sigma_{\varepsilon}^2 / \sigma_{\tilde{e}}^2$, and so, all that must be shown is that $\sigma_{\tilde{e}}^2 > \sigma_{\varepsilon}^2$. If we substitute for θ in Equation (5), we obtain:

$$\sigma_{\tilde{e}}^2\left(1 + \frac{\sigma_{\varepsilon}^4}{\sigma_{\tilde{e}}^4}\right) = \sigma_{\upsilon}^2 + 2\sigma_{\varepsilon}^2 \Leftrightarrow \sigma_{\tilde{e}}^4 - \sigma_{\tilde{e}}^2\left(\sigma_{\upsilon}^2 + 2\sigma_{\varepsilon}^2\right) + \sigma_{\varepsilon}^4 = 0$$

The solution of this quadratic equation is given by

$$\sigma_{\tilde{e}}^2 = \frac{\sigma_\upsilon^2 + 2\sigma_\varepsilon^2 \pm \sqrt{\left(\sigma_\upsilon^2 + 2\sigma_\varepsilon^2\right)^2 - 4\sigma_\varepsilon^4}}{2}$$

To ensure real solutions, we must have that

$$\left(\sigma_\upsilon^2 + 2\sigma_\varepsilon^2\right)^2 - 4\sigma_\varepsilon^4 \geq 0 \Leftrightarrow \sigma_\upsilon^2 + 2\sigma_\varepsilon^2 \geq 2\sigma_\varepsilon^2$$

which is trivially satisfied. The solution for $(\sigma_{\tilde{e}}^2, \theta)$ is not unique, suggesting that ζ_{it} can be represented by two different MA processes, both of which imply the same behavior for ζ_{it}. Without a loss of generality, we consider the larger of the two solutions for $\sigma_{\tilde{\varepsilon}}^2$ and obtain:

$$\sigma_{\tilde{e}}^2 \geq \frac{\sigma_\upsilon^2 + 2\sigma_\varepsilon^2}{2} \geq \sigma_\varepsilon^2$$

If we rule out knife-edge cases, then we have $\theta < 1$.

Appendix B

Details on the Bootstrap Procedure

To calculate the standard errors in Tables 6 and 7, we use a block bootstrap procedure. We treat each individual as a sampling unit to account for correlation in observations across time within individuals. Our procedure resamples the data 100 times. Each resampling is drawn with replacement, and the resampled data set has the same sample size as the original data set. For each resampled data set, we first estimate our main specification based on Equation (2). Using the estimated coefficients, we then compute the implied log growth rates for each scenario and employment year (where index i captures the characteristics associated with the scenario):

$$\widehat{\Delta\ln\left(Y_{it}\right)} = \hat{\alpha} + \hat{\gamma}\widehat{\Delta\ln(Y_{i(t-1)})} + \sum_{j=1}^{J} \hat{\mu}_j M_{jit} + X'_{it}\hat{\beta} \tag{7}$$

setting the initial value $\widehat{\Delta\ln\left(Y_{i0}\right)} = 0$. In a final step, we record separately the implied differences across scenarios. For instance, the comparison of 10-year cumulative growth rates across scenarios 1 and 5 is obtained as follows:

$$\sum_{t=1}^{10} \widehat{\Delta\ln\left(Y_{1t}\right)} - \sum_{t=1}^{10} \widehat{\Delta\ln\left(Y_{5t}\right)}$$

The standard errors reported in Tables 6 and 7 are the standard deviation of the relevant object over all 100 resampled data sets. Significance levels are based on the normal distribution.

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