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Are Migrants More Skilled than Non-Migrants? Repeat, Return and Same-Employer

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422

Jennifer Hunt

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Repeat, Return and Same-Employer Migrants

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**Are Migrants More Skilled than Non-Migrants?
Repeat, Return and Same-Employer Migrants**

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I am also affiliated with the NBER, CEPR, IZA, William Davidson Institute, DIW-Berlin, CIREQ and CIRANO. From 1st June 2004 I will be on the faculty of the Economics Department at McGill University (probable email: jennifer.hunt@mcgill.ca). This paper has been prepared for the 2004 Innis Lecture at the meetings of the Canadian Economic Association.

I examine the determinants of inter-state migration of adults within western Germany, using the German Socio-Economic Panel from 1984-2000. I highlight the prevalence and distinctive characteristics of migrants who do not change employers. Same-employer migrants represent 25% of all migrants, and have higher education and pre-move wages than non-migrants. Conditional on age, same-employer migrants are therefore more skilled than non-migrants. By contrast, although other migrants have higher education than non-migrants, they do not have higher pre-move wages. Furthermore, they have in their ranks disproportionate numbers of the non-employed, unemployed and recently laid off. It therefore seems inappropriate to characterize them as more skilled than non-migrants. The results for same-employer migrants indicate that skilled workers have a low-cost migration avenue that has not been considered in the previous literature. I also analyze the relation between repeat and return migration and distinguish between short and long-distance migration. I confirm that long-distance migrants are more skilled than short-distance migrants, as predicted by theory, and I show that return migrants are a mix of successes and failures.

In this paper I contribute to a literature assessing whether migrants are more skilled than non-migrants. The question is perceived as important both by sending regions, which fear a “brain drain”, and receiving regions, which believe their economy will benefit more from skilled than unskilled workers. Theoretically, the quality of migrants depends not only on relative migration benefits for the skilled and unskilled, influenced by relative inequality across regions, but also on the relation between skill and moving costs.¹

Empirical papers have often found that migrants are more educated than non-migrants in the source region, and earned higher wages than non-migrants prior to moving. Examples include Borjas, Bronars and Trejo (1993) for internal U.S. migration, and Hunt (2002) for migration from eastern to western Germany. Chiquiar and Hanson (2002) find Mexican emigrants to the United States to be more educated than Mexican non-emigrants, despite the fact that the United States-Mexico wage gap is higher for the unskilled. They believe their results point to lower moving costs for the more skilled, a point also argued by Chiswick (2000). On the other hand, it is also known that the unemployed, the laid-off, and the non-employed are more likely to emigrate than the employed. If the unemployed and laid-off are observably or unobservably less skilled than the employed, the results overall point to migrants having more heterogeneous skills than non-migrants.²

I replicate these results in my analysis of the determinants of migration by adults between federal states in western Germany. However, I show that the results indicating that many migrants are more skilled than non-migrants are strongly influenced by a group not previously analyzed: migrants who move without changing employer. Using the German Socio-Economic Panel for 1984-2000, I show that same-employer migrants represent 25% of all migrants. Same-employer migrants have

¹ Borjas (1987, 1999), Chiswick (2000).

² See also Goss and Schoening (1984), Gregg, Machin and Manning (1997), Hughes and McCormick (1981). Goss and Schoening (1984) find that annual earnings have an insignificant effect on U.S. internal migration.

higher education and higher pre-migration earnings than non-migrants. By definition, these migrants were employed before moving, and are therefore clearly more skilled than non-migrants, conditional on age. By contrast, although other migrants have higher education than non-migrants, they do not have significantly higher pre-move wages. Furthermore, they have in their ranks disproportionate numbers of the non-employed, unemployed and recently laid off. It therefore seems inappropriate to characterize them as more skilled than non-migrants.

The results for same-employer migrants indicate that skilled workers have a low-cost migration avenue that has not been considered in the previous literature. A company might request an individual to transfer, or might solicit applications for transfers in particular occupations, or might entertain a request instigated by an individual. Whichever the case, a company will most commonly want to transfer workers with skills not available locally at the destination site, which implies transferred workers will be skilled, particularly with regard to company-specific skills. Migration with the same employer is a low cost and low risk move: a new job is already associated with the move and need not be searched for, the conditions of the new job are likely to be known and similar to the old job, and skills in the old job are likely to be fully rewarded in the new job.³

Same-employer migrants could be responsible for the observed favorable characteristics of migrants in more situations than the one examined here, including migration between rich countries. They are likely to be less prevalent among migrants from poor to rich countries, however, which might appear to suggest that such migrants could be neutrally or negatively self-selected.

Distinguishing between return migrants and “new” migrants, as well as between short and long-distance moves is informative in this regard. Among migrants who are not same-employer migrants, “new” migrants moving to a non-contiguous state could arguably be characterized as more skilled than

³ Transferred workers are likely to work for large companies, which are known to pay higher wages. This may reflect unobservably higher skills of the workers or a premium with respect to the market.

non-migrants, conditional on age, while this characterization seems inappropriate for return migrants and “new” migrants moving to a contiguous state. Conditional on education, the last group of migrants is negatively self-selected. The higher skills of long-distance migrants compared to short-distance migrants is consistent with the theoretical prediction that migrants will be more positively self-selected when moving costs are high. This is the first empirical confirmation of the theory of which I am aware. The result suggests that migrants from poor countries to distant rich countries could be positively self-selected. Conversely, migrants from poor countries to neighboring rich countries need not be positively self-selected, which may go some way towards explaining the Yashiv (2004) finding of low skills among Palestinians in Israel.

Return migrants who are not same-employer migrants are a heterogeneous group of failures and successes. Laid-off males are very likely to undertake return migration, and return migration accounts for half of the positive influence of layoffs on overall migration. This is consistent with the theory of Borjas and Bratsberg (1996), that some return migrants are marginal migrants whose experience in the destination was worse than expected. In a study of U.S. internal migration, DaVanzo (1983) also found evidence for this, as return migrants were more likely to have made their initial move when unemployed, and possibly under pressure to make a quick decision.⁴ On the other hand, I find that young people with high earnings are also likely to return home. This group has likely undertaken migration as part of a lifecycle plan to raise earnings at home and consume at home, as postulated by Borjas and Bratsberg (1996) and Dustmann (2001,2003).⁵

⁴ Constant and Massey (2002, 2003) and Constant and Zimmermann (2003) find that employed foreigners in Germany are much less likely to return to their home country than their non-employed fellow foreigners.

⁵ Constant and Massey (2002, 2003), who do not split their sample by age, find that foreigners' earnings in Germany play no role in their decision to return home. Massey and Espinosa (1997) examine return migration in a non-random sample of Mexicans in the United States.

Finally, I note the link between repeat and return migration. At least 32% of all migrants in my sample have moved previously, but more than two-thirds of repeat migrants are returning to a state in which they have previously lived.⁶ I show that two-thirds of the positive effect on migration of previous migration comes through its influence on return moves. Earlier papers often attributed the effect of previous migration to an unobserved propensity to move: my results show that the contribution of such a factor will be overestimated if those at risk of a return move are not categorized separately.⁷

Theory

Chiswick (2000) contains a simple exposition of the human capital investment model of migration, whose origins are in Sjaastad (1962). An individual computes the gross benefit to migration, based on the wage difference in the source and the destination, and compares it with the cost of migration, which could include both foregone earnings and direct costs, such as transport. In the simplest version of the model, the return to skill is the same in the source and destination. If direct costs are equal for all potential migrants, migrants will tend to be those who are better paid, and therefore more skilled, since the return to migration will be higher for such individuals, all else equal. The higher the direct cost, the more positively self-selected the migrants will be. Hence, migrants who move a long distance will be more positively self-selected than migrants who move a short distance. If all moving costs are proportional to the wage, migrants will be similar to non-migrants.

⁶ Kennan and Walker (2003) make a similar observation based on the NLSY.

⁷ DaVanzo (1978) and Goss and Schoening (1984) are among previous papers to have found a large effect of a prior move on the probability of migration. Constant and Zimmermann (2003) and Massey and Espinosa (1997) study the probability of repeated round-trips between two countries. DaVanzo (1983) distinguishes return and onward migration in a sample where all have moved once.

Chiswick's model also shows simply the result developed in Borjas (1987), based on the Roy model, that relative returns to skills (inequality) in the source and destination are important.⁸ If the return to skill is higher in the destination, this factor alone implies a higher return to migration for the more skilled, and hence more positively self-selected migrants. If the return to skill is lower in the destination, this will tend to offset, and possibly even reverse, the positive self-selection of migrants.

In this paper I discuss the self-selection of immigrants as viewed from the source: I consider a migrant to be positively self-selected if he or she is more skilled than non-migrants in the source region, as measured by personal and labor market characteristics, and the wage. Borjas (1987) shows, however, that it is possible for a migrant to be paid less than non-migrants in the source, yet to earn more than natives in the destination ("refugee sorting"). The converse is not possible: migrants who earned more than non-migrants in the source will perform well in the destination.

Data

The sample from the German Socio-Economic Panel (GSOEP) includes both workers and non-workers, aged 18-65, and spans the years 1984-2000.⁹ I focus on inter-state moves of residents of western Germany, since east-to-west moves may have somewhat different determinants.¹⁰ I consider people under thirty moving with their parents to be non-migrants ("stayers"). People moving abroad are recorded as missing observations. The small state of Saarland is not identified separately in the GSOEP, and is treated as part of Rheinland-Pfalz (Rhineland-Palatinate).¹¹

⁸ See also Borjas (1999).

⁹ See SOEP Group (2001) for details on the data. Foreigners are over-sampled.

¹⁰ East-to-west migration is examined in Burda and Hunt (2001) and Hunt (2002).

¹¹ Ham et al. (2003) measure migration ideally, by measuring who moves beyond a radius of a certain distance.

In each survey, respondents are asked explicitly whether they had changed employer since the beginning of the previous calendar year, and the date of and reason for any change. By combining this information with the date of the previous interview, I can establish whether a respondent had changed employer between interviews, and whether the change had been involuntary (which I term a layoff). In the interpretation of the results, I assume that layoffs occurred before the move, whose date is bounded only by the interview dates.

I construct a variable indicating whether the respondent has moved in the previous ten years (or since their entry to the survey, if less than ten years previously), and for migrants, I check whether they have lived in their destination state in the previous ten years (or since entry to the survey). These variables will be underestimated in the early part of the survey, meaning early return migrants will be misclassified, and the effect of a previous move will be biased downward.¹² I construct an hourly wage rate from monthly labor earnings divided by 4.33 times usual weekly hours. I do not drop observations with missing layoff information, since this leads to a disproportionate loss of migrants, but instead include a dummy for having missing layoff information.¹³

Model

I begin by estimating probits for the probability that a respondent i changes state of residence r between interviews at t and $t+1$, $P(\text{move}_{t,t+1})$, as a function of characteristics before the move, X_{irt} , whether the respondent was laid off between the interviews, $\text{Laid-off}_{irt,t+1}$, whether the respondent

¹² I compute these variables based on the periods in which I observe respondents in the sample, to avoid having missing values for all who temporarily left the sample.

¹³ Dropping observations with missing wages reduces the sample considerably, but does not change the overall migration rate. The survey will underestimate migration, since migrants are disproportionately lost by the survey.

had migrated in the previous decade, $Move_{ir-t}$, and whether the respondent had completed an educational qualification between the interviews, $Edu_{irt,t+1}$:

$$P(Move_{irt,t+1}) = X_{irt}\beta_1 + \beta_2 Laid-off_{irt,t+1} + \beta_3 Move_{ir-t} + \beta_4 Edu_{irt,t+1} + \delta_r + \gamma_t + \varepsilon_{irt}.$$

I also control for year dummies γ_t and state dummies δ_r . I report marginal effects, and t-statistics adjusted to reflect repeated observations on individuals. I include information on the acquisition of an educational qualification, since many people move after completing a qualification, either to continue their education or take a job.¹⁴

I subsequently estimate multinomial logits based on this specification. I first divide migrants into three mutually exclusive groups: same-employer migrants, return migrants (who were not same-employer migrants), and others. I then divide the “others” group into those who moved to a neighboring state, and those who moved farther. This gives me four groups of migrants and the group of stayers as categories for the multinomial logit. I report odds ratios (exponentiated coefficients) and t-statistics adjusted to reflect repeated observations on individuals.

Same-employer migrants had been working in year t , were not laid off between interviews, and with very few exceptions did not report being unemployed in year t . Therefore, the coefficients on these variables for same-employment migration are constrained to be zero (with an odds ratio of one). When the sample is split by age or sex, the small number with an increase in education requires this coefficient too to be set to zero for same-employer migration. Since there cannot be any return migrants in the first year, the coefficient on the 1984 year dummy is constrained to be zero for return migration.

¹⁴ Since most moves in a period where an educational qualification was obtained are likely to take place after it is obtained, I control for educational qualifications in $t+1$ rather than t to avoid characterizing migrating graduates as having lower education than is in fact the case.

Descriptive Statistics

Column 1 of Table 1 shows that there were 829 moves in a sample with 114,946 person-year observations, or 0.72% per year.¹⁵ The subsequent columns indicate that the migration rate is twice as high among those under thirty compared to those aged 30-49 (columns 2 and 3), and more than twice as high among those laid off as those not laid off (columns 6 and 7). Men and women have similar migration rates (columns 4 and 5).

In this table I distinguish between same-employer migrants to a new state, and same-employer return migrants, who represent 20% and 5% of moves respectively (column 1).¹⁶ Other return migrants represent 17% of migrants, 30% of migrants moved to a “new” neighboring state, and 27% moved to a “new” non-contiguous state. A majority of same-employer moves are for people aged 30-49, for whom they represent in total 34% of moves (column 3), consistent with the notion that these are transferred workers who have accumulated firm-specific knowledge. Most of these moves are made by men, for whom they represent 34% of moves, rather than women, for whom they represent only 18% of moves (columns 4 and 5).

Table 2 shows the means of the main variables of interest by migrant status, again distinguishing between the two types of same-employer move. The most striking statistic of the table is the share of migrants who had completed tertiary education (university or technical higher education - Fachhochschule). While only 9% of stayers had completed tertiary education (column 1), 27% of movers had done so (column 2), while 34% of same-employer movers to a new state and fully 44% of return same-employer movers (columns 6 and 7) had done so. Same-employer migrants are slightly older than other migrants, and earned more in year t . Individuals laid off between interviews are overrepresented among movers, particularly among return movers (column 5), where they represent

¹⁵ The weighted proportion is 0.90%.

¹⁶ The weighted proportions are 20% and 8% respectively.

12% of the group. 32% of migrants (column 2) compared to only 3% of non-migrants (column 1) had moved in the previous ten years, but among non-return migrants, this share is never higher than 15%. A calculation not shown indicates that more than two-thirds of previous movers are return movers. Means of other variables used in the regressions are shown in the Appendix Table.

The two groups of same-employer migrants have similar characteristics, and in multinomial logit analysis the hypothesis that their coefficients were the same could not be rejected. I therefore pool these two groups in the multinomial results reported below.

Results – Probits

The columns of Table 3 represent a series of probits for the probability of migration, with some variations in the definition of migration. All specifications include controls for sex, marital status, the presence of a child eleven or younger, the interaction of marital status and child with sex, and dummies for foreign nationality, age, year and state, in addition to those shown in the table. The first five columns use the migration definition described in the data section. In column 1 I add education dummies to the basic set of controls. The omitted category is apprenticeship, the most common category. The dummies for the less-skilled “general education” categories have insignificant negative coefficients, as does the dummy for civil service training. However, individuals with vocational training, which typically follows an apprenticeship, are a significant 0.2% more likely to migrate, while those with tertiary education are 1.2% more likely to migrate, compared to a mean of only 0.7%. These results suggest that migrants are more skilled than non-migrants. Those that acquired an educational qualification were 0.27% more likely to migrate.

In column 2, rather than using education as a proxy for skill, I use the wage in the initial year. This is naturally only valid for those who are working. For the employed, the wage may be the best

measure of skills, since it should reflect both observed characteristics and unobserved characteristics such as motivation. Those who are not working are 0.29% more likely to move than those who are working, while amongst those who are working, those with a higher wage are more likely to migrate, indicating that the more skilled workers migrate. The mean hourly wage in the sample, as well as the mean wage of those with apprenticeship education, is DM20 (standard deviation 23), while for tertiary graduates it is DM31. The coefficient of 0.0175 indicates that were the hourly wage to rise by DM11 per hour, the probability of migration would rise by 0.2%, only 16% of the magnitude of the university effect in column 1.

In column 3 I use labor force status only as a proxy for skill (although the labor force coefficients do not prove very sensitive to the presence of other skill covariates). Similarly to column 2, those not working are 0.28% more likely to migrate than the employed. Those registered as unemployed at the time of the initial interview are 0.17% more likely to migrate than others who were not working, meaning they are $0.28+0.17=0.45\%$ more likely to move than the employed, about half the mean migration rate. Those laid off between interviews (who would typically have been employed at the time of the initial interview), are 0.58% more likely to emigrate, about half the size of the university effect.

In column 4 I include education, labor force status and wage, as well as an indicator for working part-time (whose coefficient is insignificant). The main change when the skill indicators are entered simultaneously is that the coefficient on the wage falls to two-thirds of its column 2 level. In column 5 I add the covariate indicating whether the individual had moved in the previous ten years. This is very significant and raises the R^2 of the regression considerably: a previous move raises the probability of a subsequent move by 2.7%, the largest effect identified. The only other coefficient to change appreciably when this covariate is added is the coefficient on university, which falls to two

thirds of its column 4 level: part of the reason that the university-educated are more likely to move is that they are more likely to have moved in the past.

In columns 6-7 I recode the moves of same-employer migrants to be stays, to assess the influence of this group on the results. I present results from the specifications corresponding to columns 2 and 5. Removing the same-employer migrants in column 6 causes the coefficient on the wage to fall to less than 40% of its previous size, and increases the standard error. The coefficient is thus not close to significant. If the wage is used as a summary measure of the skill of employed potential migrants, therefore, migrants who are not same-employer migrants are not positively self-selected.

In column 7 I repeat the specification with all covariates including previous move. The coefficient on the wage is here only a quarter of its column 5 counterpart and insignificant, and the coefficient on university falls to almost half its column 5 level. The coefficient on vocational training also falls, while the coefficients on not working and being laid off rise (since migrants who by definition were employed and not laid off have been removed). This reinforces the impression that migrants who are not same-employer migrants are not more skilled than non-migrants. The coefficient on a previous move falls somewhat in column 7 compared to column 5, suggesting that some of its effect comes from identifying people likely to be mobile same-employer movers.

In column 8, compared to the standard migration definition, I recode moves of those returning to a state they have lived in the previous decade, including same-employer movers, to be stays. I present the specification corresponding to column 5. The coefficient which changes the most when returners are recoded is that on a previous move: it falls to an effect of 0.85%, only 30% as large as the column 5 effect. Thus, most of the effect of a previous move works through the fact that previous movers are at risk to return home. Nevertheless, the remaining effect is still quantitatively important,

and with university education the strongest predictor of migration.¹⁷ There is also a large decrease in the coefficient on a layoff: it falls from 0.43% in column 5 to 0.22% in column 8, indicating that layoffs have a strong influence on return migration. The coefficient on not working falls when returns are coded as stays.

Results – Multinomial Logits

In Table 4 I present the results of the multinomial equivalents of columns 2 and 5 in Table 3.¹⁸ The reference category is non-migration. In columns 1-4 I focus on the wage as a summary of skill. As expected based on the results of the probits, same-employer migrants have significantly higher pre-move wages than stayers: the coefficient indicates that a one DM increase in the hourly wage would increase the relative migration probability by 0.4%, and hence a DM11 increase would increase the migration probability by 4.5%. Long-distance migrants are similarly positively and statistically significantly self-selected. Return migrants have an imprecisely estimated coefficient of zero, while migrants to a neighboring state appear to have lower wages than stayers, but the coefficient is not statistically significant. Compared to the employed, the non-employed have approximately double the relative probability of each type of migration.

In columns 5-8 I include all the skill measures. All migrants are positively self-selected on their education: having tertiary education more than triples the probability of being in one of the migrant categories compared to being a non-migrant (the coefficients are not significantly different across columns). The wage coefficients generally remain unchanged, but the coefficient for short-distance moves falls sufficiently to become significantly negative: conditional on education, the employed

¹⁷ In results not shown, if both same-employer movers and returns are coded as stays, the coefficient on a previous move falls to 0.0046 with a t-statistic of 6.0.

¹⁸ I choose not to control for a previous move, since the coefficient for the same-employer move will be difficult to interpret, because some of these movers are return movers.

members of this group are negatively self-selected (column 5). This is the wage effect least dwarfed by the education effect: a DM1 rise in the wage reduces the relative probability of this type of migration by 2.5%, and a DM11 rise reduces the probability by 27%.

The weak effects of wages relative to education could indicate that the importance of education goes beyond its relation with labor market productivity. An alternative explanation, particularly for short-distant migration for which the wage is significantly negative conditional on education, is that some people migrate because they are underpaid in their current job. This in turn could be a manifestation of Borjas's "refugee sorting".¹⁹

As expected based on the probits, the magnitude of the layoff effect is greatest for the return migration category. The relative probability of return migration is multiplied by almost six for a person who is laid off (column 7), compared to a multiple of two (and insignificant) for a short-distance move (column 5), and a multiple of 3.5 for a long-distance move (column 6). The effects are significantly different for return and short-distance migration.

Compared to other non-employed, the unemployed are significantly more likely to make a move to a neighboring state only. The total effect of being unemployed is the product of the odds ratios for non-employment and unemployment: compared to the employed, the unemployed are four times (2.40×1.69 , column 5) more likely to move to a neighboring state than to stay, which is larger than the layoff effect of 1.9, though insignificantly so. The effect of a layoff and the total effect of unemployment are similar for long-distance moves (odds ratio of about 3.5, column 6), but the layoff odds ratio is larger for return moves (5.7 vs 3.2 in column 7), though insignificantly so. It is possible that when workers who have the possibility of returning home are laid off, they return immediately

¹⁹ Hunt (2002) finds a similarly weak effect of the wage compared to education for east-to-west migration.

without waiting to register for unemployment benefits. Those who remain non-employed long enough to be picked up as such by the survey may have made their decision not to return.

The part-time employed and the newly qualified tend to make short-distance moves: the odds ratios associated with this category are about two and significant, whereas the associated coefficients are insignificant for the other moving categories.

Table 5 presents the multinomial logit specification of Table 4 columns 5-8 with the sample split by sex. For women, only same-employer migrants have relatively higher wages than stayers (column 8), whereas for men both long-distance movers and same-employer movers have higher wages than stayers (columns 2 and 4). Also, while male short-distance moves have an almost significant negative coefficient on the wage (column 1), the women's coefficient is closer to zero and far from significant (column 5). These results suggest the possibility that women who are not moving with the same employer are often tied movers.²⁰

While for men the effect of a university education is strongest for same-employer moves (column 4, odds ratio of 5.5), for women, although the same-employer effect is positive and significant, the odds ratio of 2.4 is the smallest of the moving categories (column 8). The differences between the university coefficients across categories are not statistically significant for men or women, however. The most common and most over-represented occupation for female same-employer migrants is “accountants, cashiers and related”, who are 19% of same-employer migrants compared to 9% in the general female workforce. For male same-employer migrants the most common and most over-represented occupation is “architects and engineers”, who are 17% of the same-employer migrants compared to 4% in the general male workforce.

²⁰ Interactions of the wage with sex, spouse, and sex times spouse all had insignificant coefficients, however. It is difficult to estimate the logit on single women only, since the number of moves per category becomes small.

Also of great interest in Table 5 is the result for men that those laid off are twelve times more likely than others to return home relative to staying (column 3). This coefficient is significantly larger than for the other migration categories. The result suggests that men who migrate and fail return home. No such pattern is seen for women, for whom the layoff effect operates exclusively through long-distance moves.

In Table 6 I present the same specifications for a sample split by age. I focus on young people (18-29) and prime-aged people (30-49). Layoffs and the wage have different patterns by age. Layoffs push both the young and prime-aged to return home (columns 3 and 7 indicate that a layoff increases the relative probability of such a move by 4.5-5 times), but have much stronger effects on other moves for the prime-aged than for the young (columns 1 and 5, 2 and 6). The association between same-employer moves and the wage does not differ by age, but while long-distance migrants are positively self-selected on the wage among the prime-aged, it is rather return movers who are positively self-selected in the case of the young. The unreported coefficients from the specification with only the wage as a skill measure show the wage coefficients to be similar to the reported coefficients, with the coefficient for young returners significant.

The stronger effect of layoffs on older people confirms the result of Hunt (2002) for east-to-west migration. However, for east-to-west migration, younger individuals were found to be more positively self-selected than older individuals based on the wage.

Conclusion

Skilled migrants have a low-cost avenue for migration that has not previously been considered: transferring within their company. This contributes to the positive self-selection of migrants, particularly for internal migration, or migration between rich countries. I show that this avenue is

quantitatively important for German internal migration, where a quarter of inter-state migrants move with the same employer, and account for the overall positive characteristics of migrants compared to non-migrants.

Migrants who are not same-employer migrants cannot be considered generally more skilled than non-migrants. However, it is important to distinguish further between return and “new” migrants, and short and long-distance migrants. Return migrants are a heterogeneous group of failures and successes, in keeping with theory predicting that some return migrants had worse realized outcomes of their initial migration than they had expected, while some had always intended to return. I also show that repeat migration is primarily accounted for by return migration, and that return migrants must be classified separately when associating the effect of previous migration with an unobserved propensity to move.

Among “new” migrants, long-distance migrants are positively self-selected, while short-distance migrants are not. This is the first confirmation of the theory predicting that higher moving costs increase the positive self-selection of migrants. The result suggests that migrants from poor countries to distant rich countries are likely to be positively self-selected, while migrants from poor countries to neighboring rich countries need not be.

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Table 1: Frequency and Types of Inter-State Year-to-Year Migration

	All	18-29	30-49	Men	Women	Not laid off	Laid off
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sample size	114,946	33,238	51,224	57,239	57,707	107,139	2,526
% migrated	0.72%	1.33%	0.63%	0.72%	0.73%	0.65%	1.58%
Types of migration:							
New neighbor state, not with same employer	248 (30%)	148 (33%)	80 (25%)	108 (26%)	140 (33%)	199 (28%)	10 (25%)
New non-neighbor state, not with same employer	227 (27%)	115 (26%)	93 (29%)	100 (24%)	127 (30%)	180 (26%)	15 (37%)
Return state, not with same employer	142 (17%)	92 (21%)	42 (13%)	62 (15%)	80 (19%)	109 (16%)	15 (37%)
New state, same employer	169 (20%)	74 (17%)	84 (26%)	112 (27%)	57 (14%)	169 (24%)	0
Return state, same employer	43 (5%)	14 (3%)	26 (8%)	28 (7%)	15 (4%)	43 (6%)	0
Number of moves	829 (100%)	443 (100%)	325 (100%)	410 (100%)	419 (100%)	700 (100%)	40 (100%)

Notes: Unweighted means for individuals aged 18-65. “Laid off” means that in the second year of the pair the individual reported having been laid off since the first interview. The sample includes some individuals whose layoff status is unknown.

Table 2: Means by Migration Status and Type of Migration

	Stayers	Migrants					
		All	Neighbor state	Distant state	Return state	Same employer (new state)	Same employer (return state)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sex	0.50	0.51	0.56	0.56	0.56	0.34	0.35
Age	39.4 (13)	31.8 (10)	30.9 (10)	31.7 (11)	30.3 (10)	33.7 (10)	35.7 (9)
University	0.09	0.27	0.26	0.26	0.20	0.34	0.44
Not working	0.35	0.38	0.51	0.48	0.49	0.05	0.02
Unemployed	0.06	0.10	0.15	0.11	0.13	0.01	0.02
Hourly wage (if working)	20.4 (23)	22.1 (33)	16.6 (8)	24.7 (53)	17.9 (8)	25.0 (35)	26.4 (15)
Laid off (if non-missing)	0.02	0.05	0.05	0.08	0.12	0	0
Previous move	0.03	0.32	0.15	0.11	1	0.14	1
Observations	114,117	829	248	227	142	169	43

Notes. Whether the individual was working or registered as unemployed, and the wage of the employed refer to the initial year of the pair. Education is measured in the second year of the pair, while “laid off” refers to those reporting a layoff between the two interviews. The standard deviations of wages and age are in parentheses. Wages are measured in 1991 DM.

Table 3: Sensitivity of Determinants of Migration to Same-Employer Migration and Return Migration

	Standard migration definition					Same-employer migrants are stayers		Returners are stayers
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
German general edu	-0.0008 (-1.5)	--	--	-0.0012 (-2.5)	-0.0011 (-2.7)	--	-0.0008 (-3.1)	-0.0010 (-2.4)
Foreign general edu	-0.0006 (-0.6)	--	--	-0.0008 (-0.9)	-0.0007 (-1.1)	--	-0.0003 (-0.6)	-0.0007 (-0.9)
University	0.0123 (12.0)	--	--	0.0117 (11.9)	0.0074 (10.6)	--	0.0040 (7.9)	0.0083 (11.2)
Vocational edu	0.0021 (3.3)	--	--	0.0019 (3.1)	0.0016 (3.2)	--	0.0010 (2.7)	0.0017 (3.4)
Civil service training	-0.0013 (-1.0)	--	--	-0.0009 (-0.7)	-0.0006 (-0.6)	--	-0.0008 (-1.0)	-0.0007 (-0.7)
Education increase	0.0027 (2.7)	--	--	0.0026 (2.8)	0.0030 (3.3)	--	0.0021 (3.2)	0.0032 (3.5)
Not working	--	0.0029 (5.9)	0.0028 (5.3)	0.0033 (6.0)	0.0030 (6.3)	0.0038 (8.9)	0.0041 (9.7)	0.0023 (4.9)
Work part-time	--	--	--	0.0006 (0.7)	0.0010 (1.5)	--	0.0015 (2.5)	0.0009 (1.4)
Unemployed	--	--	0.0017 (2.0)	0.0016 (2.1)	0.0014 (2.1)	--	0.0009 (2.0)	0.0013 (2.0)
Laid off	--	--	0.0058 (4.3)	0.0061 (4.8)	0.0043 (4.0)	--	0.0057 (6.3)	0.0022 (2.1)
Wage*working *1000	--	0.0175 (4.0)	--	0.0111 (2.0)	0.0099 (2.0)	0.0066 (1.1)	0.0024 (0.3)	0.0101 (2.2)
Previous move	--	--	--	--	0.0274 (22.5)	--	0.0187 (20.4)	0.0085 (9.8)
Pseudo R ²	0.10	0.08	0.08	0.11	0.16	0.09	0.18	0.12

Notes: Marginal effects from probits on a sample of 114,946 observations. The t-statistics in parentheses are adjusted to account for repeated observations on the same person, and are computed for the untransformed coefficients. Regressions also include sex, spouse, sex*spouse, child, sex*child, foreign and age dummies, year dummies and state dummies. Columns 3,7 and 8 also contain a dummy for missing layoff information. The omitted education is apprenticeship. In columns 6 and 7 same-employer migrants are coded as stayers rather than migrants. In column 8 all return migrants, including same-employer return migrants, are coded as stayers rather than migrants.

Table 4: Multinomial Estimation of Determinants of Moving

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Neighbor state	Distant state	Return state	Same employer	Neighbor state	Distant state	Return state	Same employer
University	--	--	--	--	3.77 (6.9)	3.87 (6.9)	3.55 (4.8)	4.33 (7.3)
Education increase	--	--	--	--	2.17 (3.2)	1.04 (0.1)	0.58 (-1.1)	1.19 (0.5)
Not working	1.98 (2.9)	2.45 (5.7)	2.38 (4.5)	1	2.40 (3.5)	3.56 (6.6)	2.80 (4.0)	1
Work part-time	--	--	--	--	2.05 (2.6)	1.23 (0.7)	0.70 (-0.7)	1.26 (0.7)
Unemployed	--	--	--	--	1.69 (2.3)	1.01 (0.0)	1.14 (0.4)	1
Laid off	--	--	--	--	1.90 (1.7)	3.47 (4.3)	5.72 (5.8)	1
Wage*working	0.982 (-1.6)	1.004 (3.6)	1.000 (-0.1)	1.004 (8.3)	0.975 (-2.4)	1.004 (2.8)	0.995 (-0.7)	1.004 (7.0)
Pseudo- R ²	0.09				0.13			

Notes: Odds ratios (exponentiated coefficients) for multinomial logits on a sample of 114,946 observations. The reference category is non-migration. The t-statistics in parentheses are adjusted to account for repeated observations on the same person, and are computed for the untransformed coefficients. Regressions also include sex, spouse, sex*spouse, child, sex*child, foreign and age dummies, year dummies and state dummies. The omitted education is apprenticeship. Columns 5-8 also contain a dummy for missing layoff information and the remaining education dummies.

Table 5: Multinomial Estimation of Determinants of Moving by Sex

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Men				Women			
	Neighbor state	Distant state	Return state	Same employer	Neighbor state	Distant state	Return state	Same employer
University	3.24 (4.0)	3.32 (4.3)	2.57 (2.3)	5.54 (7.1)	3.88 (5.3)	4.55 (5.5)	4.70 (4.4)	2.40 (2.3)
Education increase	3.03 (3.1)	1.18 (0.4)	0.86 (-0.2)	1	1.61 (1.4)	0.85 (-0.4)	0.37 (-1.3)	1
Not working	2.24 (2.1)	4.79 (5.4)	2.18 (1.9)	1	2.65 (2.8)	1.91 (1.9)	3.46 (3.5)	1
Work part-time	4.01 (3.3)	3.1 (2.2)	1.05 (0.0)	0.57 (-0.8)	1.64 (1.4)	0.80 (-0.6)	0.84 (-0.3)	1.70 (1.5)
Unemployed	1.52 (1.2)	1.26 (0.6)	0.93 (-0.1)	1	1.86 (2.1)	0.81 (-0.6)	1.32 (0.7)	1
Laid off	2.84 (2.1)	2.44 (1.9)	12.00 (7.0)	1	1.16 (0.3)	4.33 (4.0)	1.31 (0.3)	1
Wage*working	0.973 (-1.9)	1.005 (3.9)	0.993 (-0.6)	1.004 (5.4)	0.980 (-1.3)	0.982 (-1.2)	0.993 (-0.6)	1.004 (4.8)
Pseudo- R^2	0.13				0.13			
Observations	57,239				57,707			

Notes: Odds ratios (exponentiated coefficients) for multinomial logits on a sample of 114,946 observations. The reference category is non-migration. The t-statistics in parentheses are adjusted to account for repeated observations on the same person, and are computed for the untransformed coefficients. Regressions also include spouse, child, foreign and age dummies, year dummies and state dummies (Hamburg and Bremen are grouped with Niedersachsen), a dummy for missing layoff information and two additional education dummies. The omitted education is apprenticeship.

Table 6: Multinomial Estimation of Determinants of Moving by Age

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Age 18-29				Age 30-49			
	Neighbor state	Distant state	Return state	Same employer	Neighbor state	Distant state	Return state	Same employer
University	3.32 (4.1)	4.52 (4.8)	2.86 (2.6)	4.51 (4.5)	4.70 (5.0)	3.92 (5.1)	3.82 (3.2)	4.27 (5.6)
Education increase	2.18 (3.0)	0.87 (-0.4)	0.46 (-1.2)	1	2.53 (1.5)	1.12 (0.2)	1.08 (0.1)	1
Not working	1.95 (2.1)	2.26 (2.2)	3.31 (4.3)	1	2.29 (1.7)	4.77 (4.6)	2.37 (1.5)	1
Work part-time	3.08 (3.2)	1.80 (1.2)	1.00 (0.0)	0.79 (-0.4)	1.60 (1.0)	1.17 (0.3)	0.63 (-0.6)	1.57 (1.2)
Unemployed	1.46 (1.3)	1.17 (0.5)	0.65 (-0.9)	1	2.16 (1.8)	0.81 (-0.5)	1.71 (1.1)	1
Laid off	1.14 (0.3)	2.13 (1.6)	4.50 (3.8)	1	3.55 (2.5)	6.55 (5.1)	5.06 (2.9)	1
Wage*working	0.953 (-2.5)	0.977 (-1.1)	1.003 (1.9)	1.005 (4.5)	0.977 (-1.9)	1.007 (3.9)	0.975 (-1.3)	1.005 (4.1)
Pseudo- R^2	0.10				0.14			
Observations	33,238				51,224			

Notes: Odds ratios (exponentiated coefficients) for multinomial logits on a sample of 114,946 observations. The reference category is non-migration. The t-statistics in parentheses are adjusted to account for repeated observations on the same person, and are computed for the untransformed coefficients. Regressions also include sex, spouse, sex*spouse, child, sex*child, foreign and age dummies, year dummies and state dummies (Hamburg and Bremen are grouped with Niedersachsen), a dummy for missing layoff information, and two additional education dummies. The omitted education is apprenticeship.

Appendix Table: Additional Means

	Stayers	Migrants
	(1)	(2)
Spouse	0.66	0.37
Sex (female) * spouse	0.34	0.19
Child age 0-11 in household	0.31	0.25
Sex (female) * child age 0-11	0.16	0.14
Foreign	0.26	0.14
Age 18-21	0.09	0.10
Age 22-25	0.10	0.20
Age 26-29	0.10	0.24
Age 30-39	0.15	0.20
Age 40-49	0.21	0.15
Age 50-59	0.19	0.07
Age 60-65	0.15	0.05
German general schooling	0.22	0.17
Foreign general schooling	0.11	0.04
University	0.09	0.27
Apprenticeship	0.39	0.33
Vocational training	0.16	0.17
Civil service training	0.03	0.01
Education increase	0.02	0.07
Work part-time	0.08	0.07
Laid off	0.02	0.05
Layoff information missing	0.05	0.11
Wage*working	13.2 (21)	13.7 (28)

Notes: Unweighted means of 114,117 stayers and 829 migrants. Variables refer to the initial year of the pair, except education, which refers to the second year. “Laid off” refers to those reporting a layoff between the two interviews. The standard deviation of wages interacted with working is reported in parentheses. Means of other variables are shown in Table 2. Wages are in 1991 DM.