

An Occupational Flow Account for the United States, 2009 to 2025: Mobility, the Sources and Destinations of Each Occupation's Workers, and the Industry Shock of 2020 and 2021

Alessandro Conway

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Keywords: occupational mobility, labour market flows, imputation, flow account, shift-share, Current Population Survey

JEL classification: J62, J24, J21, C81

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Alessandro Conway*

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Planning requires knowing how often people change occupation and where each named occupation's workers come from and go, with a measure of precision, and no published source gives this. The usual source, the March supplement of the Current Population Survey, carries a hazard. A fifth of recent records were filled in from a donor, and these records show a change of occupation in 44 per cent of cases against 4.1 per cent among respondents, supply nearly three quarters of the changes the files record, and are almost never marked by the harmonised quality flag. I build an occupational flow account on respondents alone for 214 occupations from 2009 to 2025, with replicate standard errors and the origin by destination matrix. Across occupations it agrees with the official transfer rates of the Bureau of Labor Statistics, which show no sign of the imputed records. Applied to the trend, the account shows that the rise in measured mobility after 2009 belongs mostly or wholly to the imputed records, whose share doubled, while respondents show no rise. For each occupation it gives every inflow and outflow with its standard error. In the industry shock of 2020 and 2021, with the exposure of 94 occupational groups to industry growth as an instrument for their employment growth, most of the response the supplement registers came through fewer departures out of employment, a division shaped by the survey's reference periods, and in ordinary years the instrument has no power.

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*Working paper, not peer reviewed. Companion papers are *Distance and Occupation in Interstate Migration: A Gravity Measure for 209 Occupations in the United States, 2005 to 2024*, on interstate moves into occupations, *Realised Exits from the Labour Force by Detailed Occupation in the United States: A Measure, with Three Applications*, on the people recorded here as out of work, and *Outcomes after Publicly Funded Training: New Measures from Public WIOA Records*. The author works in workforce analytics in the private sector. The paper uses only public data and received no funding.

1 Introduction

When an employer, a state, or a workforce board plans for a shortage in an occupation, the plan rests on an assumption about how many people change occupation and where that occupation's workers come from, whether other occupations, people out of work, education, or fewer of its own holders leaving. In my own work I needed those figures for named occupations, with a sense of how precise each one is, and could not find them in published form. For the United States they come largely from a single survey question. The March supplement of the Current Population Survey asks each person about the longest job held in the previous calendar year. Set against the job held in the survey week, that answer gives a transition for every worker, and the question is the source of the official rates of occupational transfer (Wolf and Lockard, 2018) as well as of much of the research on occupational mobility.

Yet no published source sets out, for every named occupation, where the people who hold it were a year earlier and where last year's holders are now, with a measure of precision on each rate. Furthermore, the files that hold the answers contain a hazard. Roughly a fifth of the people in a recent supplement did not answer it, and the Census Bureau fills in their supplement from a donor, so the occupation they are recorded as holding last year is in fact another person's. Those records show a change of occupation in 44 per cent of cases, against 4.1 per cent among people who answered, and over 2022 to 2025 they supply 73 per cent of the changes of detailed occupation the files record, yet the quality flag distributed with the harmonised data marks none of them (Section 2).

For these reasons I build an occupational flow account on the people who answered the supplement, and on them alone. An occupation's workforce is treated as a stock with three recorded inflows (people arriving from education, from other non-employment, and from other occupations) and two recorded outflows (people now in another occupation and people not at work in the survey week). The account covers 214 occupations from 2009 to 2025, with a standard error on each rate from the 160 replicate weights the Census Bureau publishes, and the full origin by destination matrix is published with it. How the account is built, what respondents' answers actually measure, and what the account can and cannot be used for are set out in Section 3. As a check, the transfer rates of the Bureau of Labor Statistics follow the account across occupations and show no sign of the imputed records (Section 3.5).

The first question I put to the account is the trend. Measured mobility rose after 2009, but over the same years the share of imputed supplements more than doubled and the rate of occupational change on all records rose with it, by 0.30 points a year, while among respondents the rate shows no trend (Section 4.1). For each of the 214 occupations the account gives each recorded inflow and outflow with its standard error (Section 4.2). The last question is how the flows among respondents responded to the contraction and recovery of industries in 2020 and 2021, with each occupational group's exposure to national industry growth as an instrument for its employment growth (Bartik, 1991). Within the survey, the change in an occupation's workforce relative to last year's holders is the sum of three arrival rates less two departure rates, so the response divides among the routes in the same way that Blanchard and Katz (1992) divide a region's response among employment, unemployment, participation, and migration. Of the

response the supplement registers, most came through fewer departures out of employment and about a fifth through arrivals from other occupations. But the division is shaped by the survey's reference periods, and outside that episode the instrument has no power, so this application describes a single episode, and the response of occupations to demand in general lies beyond it (Section 4.3).

The account classifies an arrival by where the person was the year before, a labour force state that does not identify how the person qualified for the work. For instance, a nursing aide who completes a degree while employed arrives in nursing from another occupation, so the arrivals from education describe where entrants were, which is a different thing from the share of an occupation's supply that training provides. Section 5 gathers the limits of the account in one place, and Section 6 concludes.

2 The gap

2.1 The published figures

The official occupational separations of the Bureau of Labor Statistics combine a rate of leaving the labour force with a rate of transfer to another major group (Wolf and Lockard, 2018), and the transfer rate is modelled on ten supplements and published as a projected rate for each detailed occupation (Bureau of Labor Statistics, 2026,b). Outside the official statistics, work on gross worker flows established that net employment change is the small residual of much larger movements between employment, unemployment, and inactivity (Davis and Haltiwanger, 2014). Work on occupational mobility has documented how often people change occupation and how the rate rose over three decades (Kambourov and Manovskii, 2008), how it moves with the cycle in the monthly survey (Moscarini and Thomsson, 2007), and how unemployed workers search across occupations over the cycle (Carrillo-Tudela and Visschers, 2023), while recent work treats transitions between occupations as a network (Cheng and Park, 2020; Lin and Hung, 2022). On measurement, Mellow and Sider (1983) showed that worker and employer reports of the same job frequently disagree, and vom Lehn et al. (2022) reconcile the retrospective and linked measures of mobility in this survey. Carrillo-Tudela et al. (2025) correct measured mobility for miscoding, and Kim and Kim (2026) find that measures built from the supplement's two occupation questions drew apart from 1991 to 2020, a divergence they attribute to coders. These studies are concerned chiefly with the rate of mobility, however, and none gives a planner, for a named occupation, the realised arrivals by origin and the realised departures by destination with a standard error on each. Any figure built from the supplement also depends on how the files treat the people who did not answer it. To this literature I add a realised account on respondents for every occupation the survey can support, with a standard error on every rate. I also check the account against the official transfer rates, measure the size and the course of the imputation effect to 2025, and estimate how the flows responded to the industry shock of 2020 and 2021. In a companion paper, *Realised Exits from the Labour Force by Detailed Occupation in the United States: A Measure, with Three Applications*, I follow the people this account records as not employed in the survey week to the interview a year later and set

their rates of leaving the labour force beside the official ones.

2.2 Imputed supplements

I compute every figure in this paper on people who gave a supplement interview, because of the size of what the others would add. Among people employed in both years over 2022 to 2025, the full file records a different detailed occupation for 12.1 per cent, but the figure is 44.2 per cent among those whose supplement was imputed and 4.1 among respondents. Imputed supplements are 20.1 per cent of these people and 73.2 per cent of the changes of occupation recorded. The variable that identifies them takes two values for such records, one where part of a supplement was obtained and one where an interview was obtained and could not be used, and the rate of change is 44 and 43 per cent on the two. More than half of imputed records, 56 per cent, nevertheless show the same detailed occupation in both years, far more than donors drawn without regard to occupation would give, so the procedure evidently matches on characteristics close to the recipient's occupation. I could not find the matching variables documented in the sources I consulted, but nothing here relies on an account of them. Over the full run from 2009 the three rates are 11.0, 41.3, and 4.3 per cent, and the share of recorded changes is 67.0 per cent.

Aggregating the classification does not remove the changes these records generate. As a donor is chosen to resemble the recipient, the occupation it supplies is often a neighbour of the recipient's own, and 76 per cent of the detailed changes on imputed records stay within a major group, against 20 per cent of those respondents report. The file contains 96 records of a physical therapist becoming a registered nurse, a move that requires a separate professional licence, and every one of them has an imputed supplement, whereas of the 164 records of a licensed practical nurse becoming a registered nurse, a recognised route, 19 do.

Nor can the excess be traced to non-respondents' observed characteristics. A logit model of the respondents' rate of change on year, age band by sex, education, nativity, state, current major group, and employment status, all recorded in the monthly interview, was fitted on respondents and applied out of sample to the people whose supplement was imputed over 2022 to 2025. It gives 4.3 per cent, against the 44.2 per cent their records show. The gap is consistent with an effect of the donor of the kind Hirsch and Schumacher (2004) and Bollinger and Hirsch (2006) describe for earnings, where a donor matched on some of the recipient's characteristics, but only loosely on the variable an analysis turns on, weakens the relation between that variable and the imputed one. Here the variable matched loosely is the recipient's own current occupation.

The hazard is already documented in the literature on this survey. Kambourov and Manovskii (2013) showed that changes in the imputation procedure moved measured mobility in the March files, and vom Lehn et al. (2022) drop records with an imputed occupation, whole imputed supplements included. But the harmonised quality flag distributed with the data marks only the occupations allocated on their own, while the indicator for a whole imputed supplement is available as an unharmonised source variable, and among people aged sixteen and over the first identifies 6,759 records and the second 378,055. A check of the item flag

alone therefore finds very little. For instance, Lin and Hung (2022) report an imputation rate for occupation of between 0.13 and 2.74 per cent in the supplements from 1983 to 2017, which is the order of the item flag, and a rate of change of 8.6 per cent for 2008 to 2017 among people aged 25 to 64. On all records for the same ages in the supplements of 2009 to 2017 the rate here is 8.7 per cent, and among respondents it is 3.7, so their rate appears to include imputed supplements, although the classifications and years differ somewhat.

Whole-supplement non-response in this survey has been studied chiefly for earnings, where Bollinger et al. (2019) show that it is not ignorable in the tails of the distribution, and the restriction to respondents I make here is open to the same concern, which Section 3.4 discusses.

3 The measure

3.1 Construction

Occupation in the Current Population Survey is recorded by asking what kind of work a person does, writing down the answer, and coding the description. Its March supplement returns to the question for the previous calendar year, asking about the longest job held during that year and whether it was the same as the job already recorded for the survey week. Where it was the same, the description is not taken again and the current code stands for both years, and only a respondent who reports a different job is asked to describe it, a design in use since the early 1970s (Kambourov and Manovskii, 2013; vom Lehn et al., 2022). Thus a change of occupation is recorded only where the respondent has first said that the job was different.

The account uses the seventeen annual supplements from 2009 to 2025, as distributed by IPUMS (Flood et al., 2025), comprising 2,924,246 person records of all ages, of whom 2,231,260 are sixteen or over. Occupation is the OCC2010 variable IPUMS constructs, which maps the 2002, 2010, and 2018 Census classifications used in these files to a single set of codes. The survey's changes of classification in 2011 and 2020 therefore do not create a change of occupation within a record, although a few occupations begin, end, or change in scope at those dates. These codes fall in 94 minor and 22 major groups of the Standard Occupational Classification. A transition is the pair formed by a person's longest job of the previous calendar year and the job held in the reference week of the supplement interview, so the interval it spans varies across people, and a job held during the year that was neither the longest nor the current one is not seen. For most of the sample the interview falls in March, but the files also include the households added to the March sample for the supplement, which are interviewed in February, March, or April, so a reference to the March week below stands for the week of the supplement interview.

Not everyone in the file gave a supplement interview. A person in the monthly sample for whom no usable supplement interview was obtained is given a supplement taken from a responding donor, and the share of people aged sixteen and over treated this way rose from 9.9 per cent in 2009 to 22.2 per cent in 2020 and averages 19.7 per cent over 2022 to 2025. For such a person the current occupation comes from their own monthly interview while the occupation

last year comes from the donor's supplement, so the pair describes no one at all. I drop these records, together with the far smaller number whose supplement was answered and whose occupation last year was allocated on its own (a flag that is zero in every file from 2020, so that from then the restriction rests on supplement imputation alone). That leaves 2,438,675 records, 1,846,446 of them aged sixteen or over. Their weights are scaled to reproduce the weighted total of everyone in the same year, age band, sex, and current employment status, all of which come from the monthly interview and are known for respondents and non-respondents alike.

For each occupation and year the account records the number of people whose longest job of the previous year was in the occupation, the number arriving from education, from other non-employment, and from another occupation, and the number departing to another occupation or out of employment. It includes everyone with a civilian occupation code, among them a small number aged fifteen (0.2 per cent of holders). Arrivals from outside employment have the previous year occupation code reserved for people who did not work at all, and the stated reason for not working separates education from caring, illness, retirement, and inability to find work. An arrival from education is therefore a person who did no paid work in the previous year because they were studying. By contrast, a student who held any job during the year is recorded under that job and arrives in the account as a lateral mover or a stayer, as does anyone who left education and began work before the end of that year. Section 4.2.2 shows how much the figures depend on this. Every figure comes from a single source, and no flow is calibrated to an external total.

The 2014 supplement was fielded as a split sample, each half weighted to the full population, and I keep the three eighths file, whose questions are the ones carried forward from 2015, with its own weights and replicates.¹

Standard errors come from the 160 person replicate weights the Census Bureau constructs by successive difference replication (Fay and Train, 1995), with the multiplier of 4/160 the documentation prescribes (U.S. Census Bureau, 2024). Rates are recomputed within each replicate, as numerator and denominator move together, and the scaling of respondents' weights is repeated within each replicate as well, so the errors include its sampling variation.

3.2 Coverage and precision

Precision degrades quickly as occupations get smaller. The median standard error on the rate of departure to another occupation is 0.9 percentage points where at least three hundred prior year holders are sampled and 1.6 points between one hundred and three hundred, and I report a cell only where at least one hundred holders appear. That leaves 214 occupations reported in at least one year and 151 in at least one of the four years from 2022 to 2025, out of the 474 the survey distinguishes. The four year figures in Section 4.2 are the mean of the annual rates over the years when an occupation clears the threshold, with standard errors that take the years as

¹The companion paper on leaving work draws on the same files and defines its sample differently in several respects (which half of the 2014 file is kept, the cells within which respondents' weights are scaled, the ages and armed forces status of holders, the link across years, and the years reported). For that reason the two papers' figures for the same occupation differ slightly: registered nurses not employed, for example, are 4.6 per hundred in Table 2 and 4.7 in Table 2 of the companion paper.

independent.

3.3 What respondents' answers measure

Among respondents, the supplement's measure removes one source of false change and leaves another. The occupation code for last year is carried over from the current job unless the respondent reports that the longest job last year was a different one, so for a person who reports the same job two descriptions of it are never coded twice. A person who reports a different job, by contrast, has it described and coded afresh, including one who changed employer within an occupation, and for that person a difference between the two codes can arise from the coding itself, as Lin and Hung (2022) and Carrillo-Tudela et al. (2025) both note. Recorded change is 2.7 per cent among people who worked all fifty two weeks of the previous year and between 9 and 13 per cent among those who worked fewer, the group where changes of employer concentrate. The same design misses a change that the respondent does not report as a different job, and it records only the longest job of the year. From the joint behaviour of the retrospective and linked measures, vom Lehn et al. (2022) estimate that true annual mobility is two to three percentage points, or 60 to 70 per cent, above the retrospective rate. On their estimate the 4.1 per cent recorded here understates true mobility, and the account is best described as the moves respondents themselves identify. Kambourov and Manovskii (2013) argue in addition that the retrospective question measures mobility over a period shorter than a year, a further reason not to read the rates as annual in a strict sense.

The pattern across levels of the classification fits that reading. On the supplements from 2017 the recorded rate is 4.2 per cent at the detailed level, 3.9 across the ninety four minor groups, and 3.4 across the twenty two major groups, so 80 per cent of the changes recorded cross a major group boundary. That share should not be read as showing that reported changes are mostly changes of field, however. In the linked sample below, two independent codings of an unchanged job differ across major groups in about two thirds of the cases where they differ at all, so fresh coding of a new job within the same occupation would produce a similar share. In that light, the definition of a transfer in the official separations, a change of major group (Wolf and Lockard, 2018), covers about four fifths of the changes respondents report.

The alternative to the retrospective question is to follow a person across interviews. Under the rotation design a household returns to the sample twelve months after its first interview, so the same person appears in March of one year and March of the next, with occupation collected at the time on both occasions and an eight month gap during which no answer is carried forward. I match records on the IPUMS person identifier and keep them where sex agrees and age has advanced by no more than two years, the check IPUMS itself recommends. Linking the supplement to the March monthly file for the same month gives the same current occupation for all but 13 of 394,037 respondents, so the two files share a single coding of the present job. Table 1 sets three measures of annual change beside one another for the 117,512 respondents linked a year apart and employed on both occasions.

The first two rows compare codes assigned in separate interviews, and both sit near fifty per cent at the detailed level and near a third across major groups. The third row, by contrast,

Table 1: Three measures of the same annual transition, per cent of linked respondents whose two codes differ, 2017 to 2025

	Detailed (474)	Minor group (94)	Major group (22)
Last year's job, coded then and reported now	48.1	40.9	31.6
A year apart, two independent codings	49.2	42.0	32.7
A year apart, both codes from one interview	3.1	2.8	2.4

is the supplement's own measure on the same people, and on the supplement as a whole that measure gives 4.2, 3.9, and 3.4 per cent.

Consider only the respondents whom the supplement records as having held the same occupation in both years and who worked all fifty two weeks. For these 99,684 people the code assigned a year earlier differs in 47.7 per cent of cases at the detailed level, 40.6 across minor groups, and 31.3 across major groups. Some of that is real change the screening question missed, and some of these people did change occupation between March of the prior year and the longest job of that year. But neither can account for a third of full-year workers crossing a major group boundary, when the retrospective rate, even raised by the 60 to 70 per cent that vom Lehn et al. (2022) estimate, is under a tenth. Those authors conclude that error in the linked measure is large and has grown, and the figures here extend that result to 2025. Estimates of true mobility still differ widely: Carrillo-Tudela et al. (2025) correct the linked measure for miscoding and report annual mobility of about a fifth where the uncorrected figure is over two fifths, which is several times the retrospective rate, and nothing here decides between the two. Figure 1 sets the three figures against the movement the supplement measures on the same people.

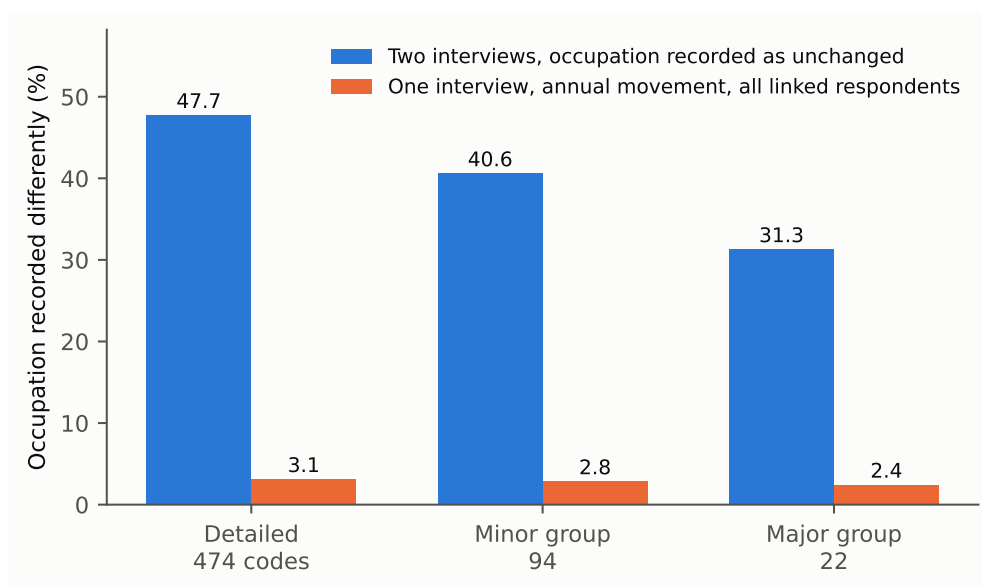


Figure 1: Two independent codings of a job recorded as unchanged disagree about half the time. The supplement's own measure records annual movement many times smaller at every level of aggregation.

Two features point to the coding as the source of the disagreement. Education, which people report afresh at each visit and which does change over a year, differs in 12.6 per cent of cases against 47.7 for occupation, and the occupational disagreement stays between 47.0

and 49.6 per cent in all nine year pairs from 2017, pandemic years included. Furthermore, the largest confusions, among them elementary against secondary teachers, supervisors against the staff they supervise, and accountants against bookkeepers, sit on boundaries in the classification where a description could be coded either way, the pattern that Carrillo-Tudela et al. (2025) find rises with the similarity of tasks.

The basic monthly files offer no cleaner measure. They carry the previous month's employer and duties forward unless the respondent reports a change (Moscarini and Thomsson, 2007), although since 2007 earlier answers may not be shown to a different member of the household without consent (IPUMS CPS, 2026), and they carry no allocation flag for occupation. In 7.7 million pairs of consecutive interviews over 2009 to 2026 where a person is employed both times, 6.7 per cent show a different detailed occupation, about twice the 3.5 per cent Moscarini and Thomsson (2007) report for men alone from 1979 to 2006, and 75 per cent of these changes cross a major group. Such a rate would imply between 44 and 56 per cent of workers changing occupation in a year, which marks it as mostly recoding. Two codings a year apart, at the fourth and eighth or the first and fifth interviews, differ in 50.1 per cent of cases over 2010 to 2026, rising from 46.5 per cent in 2010 to between 50.0 and 51.8 from 2016, in line with the growth in coding error that vom Lehn et al. (2022) and Kim and Kim (2026) describe.

Discordance concentrates within occupational families and along the supervisory boundary. For elementary against secondary teachers, discordance among unchanged jobs runs at 14.9 per hundred holders of either occupation against a recorded flow of 0.29, and for nursing aides against personal care aides at 13.3 against 0.34. As the recorded flow is built from respondents who said their job had changed, the discordance does not inflate it, but a pair of this kind is best read at a level of aggregation containing both. For registered nurses against physical therapists, by contrast, discordance is 0.08 per hundred, and among respondents the supplement records no movement between the two in the nine years.

3.4 Respondents and non-respondents

Restricting the account to respondents assumes that, within a cell of year, age, sex, and employment status, people who gave no supplement interview moved as respondents did. That assumption can be checked on current occupation, which is known for everyone. The share of the employed without a usable supplement runs from 14 to 24 per cent across major groups. In about a tenth of the larger occupations, the respondents' weighted count of an occupation's current workers differs from the count on all records by more than 5 per cent, so the published levels for an occupation are subject to an error of that order. Rates are less exposed: weights that also match the current major group leave the mobility rate at 4.1 per cent, the two whole-matrix ratios at 4.1 and 1.4, and the departures of registered nurses at 6.0 per hundred. Whether people who are hard to interview change occupation more often than respondents with the same observed characteristics cannot be checked. Differences in observed characteristics, at least, matter little, as weighting respondents by the inverse of a propensity to answer, estimated from year, age band by sex, education, nativity, state, current major group, and

employment status, moves the rate for 2022 to 2025 from 4.08 to 4.12 per cent. Dropping imputed records therefore removes a known error and, on that evidence, introduces no large one through the composition of respondents.

3.5 Agreement with the official transfer rates

Where an official figure exists, the account can be checked against it. The official transfer rates are estimated independently of this paper, although they rest on the same survey question, and so serve as an independent check on the respondent measure. To estimate them, the Bureau of Labor Statistics models transfers on ten supplements, counting a person who worked in the previous year and is now employed in a different major group, and the sample restrictions it publishes do not mention imputed supplements (Bureau of Labor Statistics, 2026). It does not name the ten years, and I take them to be 2016 to 2025. On that definition, with the transfer read as a change of major group on the harmonised codes and a universe of people now employed or out of the labour force, the transfer rate is 4.1 per cent on all records and 3.2 among respondents. Records with an imputed supplement are 20.6 per cent of the people covered and supply 37.1 per cent of the transfers. Adding the unemployed to the universe gives 3.96 and 3.09 per cent, the same ratio. Had the Bureau kept those records, its estimation data would sit about 28 per cent above the rate among respondents, although I do not reproduce the projections and the documentation does not say whether it keeps them.

The Bureau publishes a projected transfer rate for each detailed occupation (Bureau of Labor Statistics, 2026b), which allows a check that does not depend on its level. If the respondent measure is sound, the official rate should follow it across occupations, and if the imputed records entered the Bureau's estimation, occupations whose rate on all records is raised most by imputation should be the ones whose official rate sits furthest above the rate of respondents. I compute both rates for each occupation by the occupation held last year, on the wider universe above that includes the unemployed, for the 292 occupations with at least 300 respondents over 2016 to 2025, and match 247 of them to the Bureau's codes, 82.7 per cent of the weight. The Bureau's projected rates average 5.2 per cent against 3.8 on all records and 3.1 among respondents for the same occupations, a difference in level that the definitions, the projection, and the Bureau's model can produce without any help from imputation. Across occupations, however, the official rate has a weighted correlation of 0.84 with the respondent rate and 0.68 with the rate on all records. Regressed on both the respondent rate and the excess that imputation adds to it, the official rate loads on the first (0.93, standard error 0.06) and not on the second (0.02, 0.06), and as both occupation rates are measured with a reliability above 0.9, the difference does not come from noise in one of them. In other words, across occupations the respondent measure agrees with the official rates, which show no sign of the imputed records, and either the Bureau excludes them or its model, which estimates the probability of a transfer from the characteristics of workers, does not pass their occupational pattern through. Whether its level is affected cannot be settled from the published tables.

3.6 What the account can be used for

The two sides of the account do not cover the same window. An arrival from outside employment requires a whole calendar year without work, while a departure out of work requires only that a person who held the occupation at some point last year is not employed in one week of March. Over 2022 to 2025 an average of 171.0 million people a year held an occupation in the previous year and 160.9 million were employed in the March week, while arrivals from outside employment come to 4.5 million a year and holders not employed in the reference week to 14.7 million. In all but 6 of the 151 occupations reported, fewer people are at work in the occupation in March than held it during the previous year, with a median ratio of 0.95. Departures therefore exceed arrivals in nearly every occupation by construction, and the account cannot be used to tell whether an occupation is short of arrivals or losing too many people. Nor are the 14.7 million all out of the labour force: 31 per cent are unemployed in the reference week, 21 per cent retired, 5 per cent unable to work, and 41 per cent out of the labour force for other reasons, and 2 per cent are serving in the armed forces.

For any occupation with enough sampled holders the account gives the five rates and their standard errors by year, and the matrix is published with the paper as the mean annual flow over 2009 to 2025, at the three levels of the classification, each cell with its standard error and the number of people it rests on. Even pooled over the seventeen years the matrix is thin at the detailed level, where 687 of the 18,887 ordered pairs with any movement rest on ten or more people and hold 29 per cent of the movement. Across minor groups 1,029 pairs do, holding 75 per cent, and across major groups 361 of 456 hold 99 per cent, so work on individual pairs is better supported at minor group than at the detailed level.

Gross movement between occupations is also several times the change in occupation sizes it produces. To measure this, the absolute net lateral inflow of every occupation is summed and halved, so that each mover is counted once, with allowance for the upward bias that sampling gives an absolute difference. The firmest reading is across major groups on the pooled flows of 2009 to 2025, where lateral movement changes the sizes of groups by between about 6 and 8 per cent of the movement itself. At the detailed level in recent years, by contrast, the measured net change is mostly sampling noise.

4 Applications

4.1 The trend in occupational mobility

Whether Americans change occupation more often than they used to has a literature of its own (Kambourov and Manovskii, 2008), and in these files the answer turns on the imputed records. Their share has not been constant, and Figure 2 shows what that does to measured mobility. Among people employed in March who held an occupation the previous year, the share with an imputed supplement rose from 9.9 per cent in 2009 to 23.0 in 2019 and was 18.3 in 2025. Over the same years the rate of change on all records rose from 7.8 per cent to between 11.9 and 12.9 in 2018 to 2023, a trend of 0.30 points a year (0.07, with errors robust to

serial correlation), and the rate on imputed records itself rose from 37 to 45 per cent. Among respondents, by contrast, the rate moved between 3.5 and 4.8 per cent with no trend (-0.03 , standard error 0.02). How much of the increase of 3.3 points on all records between 2009 and 2025 the rising share accounts for depends on the year at which the share is held, with both rates at their observed values. Held at its 2009 level, the share leaves a trend of 0.02 (0.02) and accounts for all of the rise. Held at its 2025 level it accounts for 2.6 points, or 80 per cent, and leaves a trend of 0.07 (0.03), and held at the midpoint it accounts for 3.0 points. The rising share therefore accounts for between 80 per cent and all of the rise, depending on the base year, and the remainder comes from the rise in the rate on imputed records themselves.

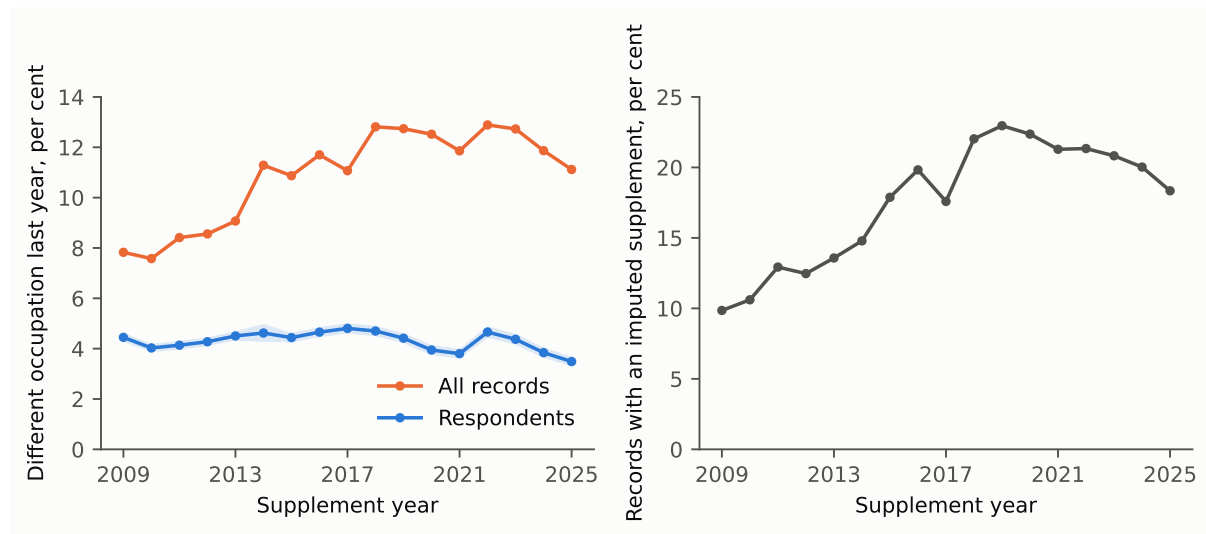


Figure 2: People employed in March with a different detailed occupation the previous year, on all records and among respondents with a 95 per cent interval (left), and the share of those records with an imputed supplement (right), by supplement year.

The same trend distorts the relation with the business cycle. Over the seventeen years the rate on all records has a correlation of -0.94 with the national unemployment rate in March, because unemployment fell over the period while imputation rose, and the respondent rate a correlation of -0.09 . Yet respondents do move with the cycle once the trends are set aside. In changes from one year to the next the respondent rate has a correlation of -0.83 with unemployment and the rate on all records -0.54 . After a linear trend is removed from each series the correlations are -0.87 and -0.79 , in line with the procyclical mobility that Moscarini and Thomsson (2007) find in the monthly survey, so only the trend on all records is spurious. Across states the relation is weak in both samples: with state and year fixed effects, a point of state unemployment goes with -0.03 points of occupational change among respondents (standard error 0.04) and -0.06 on all records (0.12).

4.2 The sources and destinations of each occupation's workers

4.2.1 Arrivals and departures

Where were an occupation's current workers a year ago, and where are last year's holders now? The account answers both questions for each of the 214 occupations it reports, with a standard error on each rate, and the published matrix gives the moves between every pair, so the same questions can be put to any of them, with the precision its size allows. Table 2 gives the account for eight occupations, averaged over 2022 to 2025 and expressed per hundred people who held the occupation the previous year, and Figure 3 shows arrivals and departures on a common scale.

Table 2: Arrivals and departures per 100 prior-year holders, average of 2022 to 2025, standard errors in parentheses

Occupation	Arrivals from			Departures to	
	Education	Other non-work	Another occupation	Another occupation	Out of work
Registered nurses	0.7 (0.2)	0.8 (0.2)	1.3 (0.2)	1.3 (0.2)	4.6 (0.4)
Elementary and middle teachers	0.5 (0.2)	1.0 (0.2)	2.1 (0.2)	2.0 (0.2)	5.2 (0.4)
Secondary teachers	0.3 (0.1)	0.5 (0.2)	2.4 (0.5)	2.5 (0.5)	4.4 (0.7)
Software developers	0.4 (0.1)	0.5 (0.1)	1.4 (0.2)	1.6 (0.3)	4.8 (0.5)
Secretaries and administrative assistants	0.4 (0.2)	1.2 (0.2)	3.2 (0.4)	2.8 (0.4)	7.2 (0.5)
Truck drivers	0.4 (0.1)	1.9 (0.2)	3.2 (0.3)	3.0 (0.3)	9.9 (0.5)
Nursing and home health aides	0.9 (0.2)	2.0 (0.3)	4.3 (0.5)	3.7 (0.4)	10.1 (0.7)
Cashiers	4.7 (0.4)	3.3 (0.3)	4.8 (0.4)	6.6 (0.4)	20.4 (0.8)

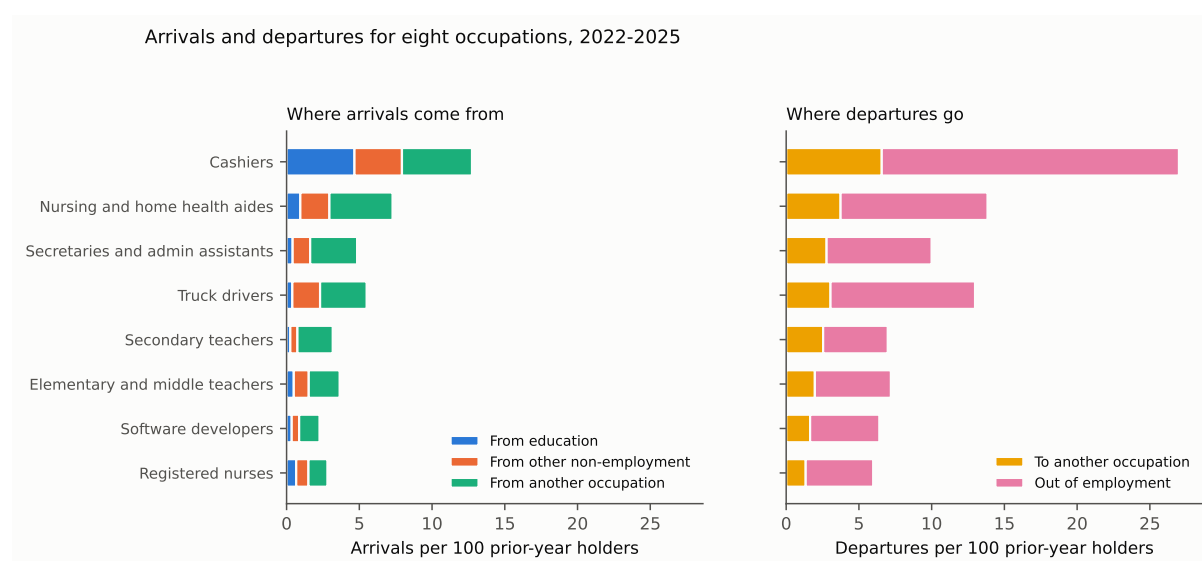


Figure 3: Arrivals and departures per 100 prior-year holders for the eight occupations, average of 2022 to 2025.

In all eight occupations more of last year's holders are out of work in the March week than are in another occupation. Of every hundred people who were registered nurses last year, 6.0 (standard error 0.4) are no longer nurses in March, 4.6 of them not employed, of whom 1.8 are retired and 1.0 unemployed. Among the health and education occupations with a million or more holders, physicians have a lower figure, at 3.2 (0.5), and nursing and home health aides a much higher one, at 13.8 (0.8). Secondary teachers at 7.0 (0.8), elementary teachers at 7.2 (0.4), and software developers at 6.4 (0.6), by contrast, cannot be distinguished from nurses, the differences being 1.0 (0.9), 1.2 (0.7), and 0.4 (0.8). Of the gap of 7.9 (0.9) between nurses and aides, about 5.5 is people not employed and the remainder is movement to other occupations.

On the arrival side, another occupation is the largest of the three origins in all eight, although for cashiers its margin over education is well within the standard error, and for registered nurses and for cashiers the two origins outside employment together exceed it. Across all the occupations reported, education is the smallest of the three inflows in 77 per cent, and twenty one record no arrival from education at all across the four years. The median occupation takes in 7.5 people from another occupation for each one recorded as arriving from education, and a quarter of occupations take in fewer than 3.2, with the twenty one that record no arrival from education placed at the top of the distribution.

Figure 4 plots every occupation reported. In 117 of the 151, or 77.5 per cent, arrivals from other occupations exceed arrivals from outside employment. The thirty four below the line include registered nurses, physicians, pharmacists, physical therapists, and licensed practical nurses, the food service and cleaning occupations, cashiers and retail salespersons, and several construction trades. Those in the first group are licensed, which would limit entry by people working elsewhere, while the others are occupations of high turnover, although I do not test either explanation.

Across the whole matrix, arrivals from another occupation run 1.4 times all arrivals from outside employment at the detailed level (standard error 0.03), 1.3 at minor group, and 1.1 (0.03) across the twenty two major groups. Against arrivals from education alone the ratio is 4.1 (0.1) at the detailed level and 3.3 (0.1) at the coarsest, and over the full run from 2009 the corresponding figures are 4.2 and 3.3. This ratio varies from year to year by more than sampling error, between 2.9 and 4.3 at major group. The ratios change little with the level of aggregation, as most of the changes respondents report cross a major group boundary, a point Section 3.3 discusses.

4.2.2 The age profile

Arrivals recorded as coming from another occupation might include graduates who held a job while studying, which would concentrate them in the youngest bands, and Figure 5 shows that in three occupations the rate is highest in the youngest band. Across the whole matrix 59 per cent of lateral arrivals are under thirty five, against 35 per cent of the employed population, although arrivals from education are younger still, with 92 per cent of them under thirty five. Lateral arrivals aged thirty five and over, few of whom can be recent graduates under another heading, come to 1.7 times all arrivals from education as the survey defines them, 1.5 at minor

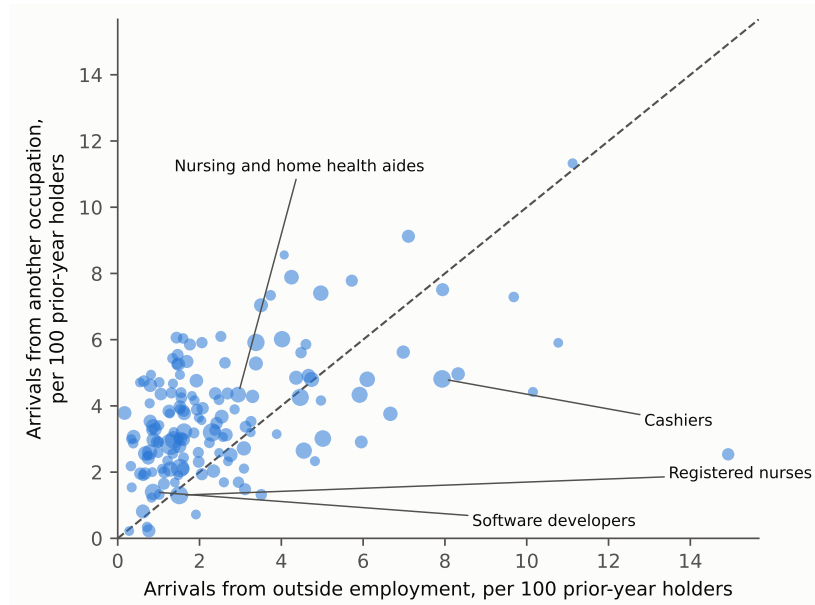


Figure 4: Arrivals from another occupation against arrivals from outside employment, one point per occupation, 2022 to 2025.

group, and 1.3 at major group.

The larger issue, however, lies on the other side of the ratio. Arrivals from education as the survey defines them number 1.57 million a year, and 51 per cent of them are under twenty, which is to say that the category is mostly teenagers taking a first job and holds few graduates. A graduate who started work during the previous year is recorded under the job held, and is a stayer if still in it. Each year 4.66 million people aged twenty four or under are employed in March after working between one and thirty nine weeks of the previous year, and 85 per cent of them are recorded in the same occupation. Not all of them left education that year, and some who did are older, so no single line identifies entrants. Counting everyone aged twenty four or under who worked one to thirty nine weeks as an arrival from education gives a ratio of lateral arrivals to arrivals from education of 0.9 where the survey's definition gives 4.1, and a ratio to all arrivals from outside employment of 0.6 where it gives 1.4. Drawing the lines at twenty two and twenty six weeks gives 1.4 and 0.8, and drawing them at twenty four and fifty one weeks gives 0.6 and 0.5. On the first of these wider definitions, lateral arrivals aged thirty five and over are 0.4 times arrivals from education, against the 1.7 above. As such, the data do not support a statement that occupations draw several times as many people from other occupations as from education, and they do not settle which of the two is larger.

4.3 The industry shock of 2020 and 2021

4.3.1 Design

How did occupations' flows move when the industries that employ them contracted and recovered in 2020 and 2021? Among respondents the flows can be set against that single large and uneven shock, and I treat the estimate as a description of that episode alone, leaving aside

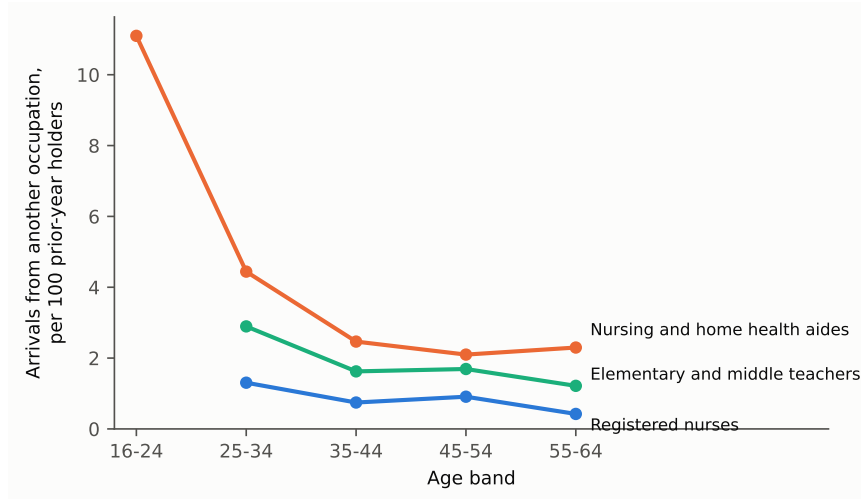


Figure 5: Arrivals from another occupation per 100 prior-year holders, by age band, 2022 to 2025.

the response of occupations to demand in general. The rates per hundred holders last year satisfy an identity: the occupation's workforce in March relative to last year's holders, less one, equals arrivals from another occupation, from education, and from other non-employment, less departures to another occupation and out of employment. As such, a regression of each of the six rates on the same measure of demand, with the same controls and instrument, divides the response of the left side exactly among the five routes.

Demand is measured as the growth in an occupation's average monthly employment over the twelve months from April to March, taken from the basic monthly files, which hold about twelve times as many person-months a year as the supplement, and growth in year t is set against the supplement of March t . Employment growth reflects supply as well as demand, so it is instrumented with a shift-share variable (Bartik, 1991): the occupation's distribution across 221 industries in the twelve months to March 2012 times the growth of each industry's national employment, with the occupation's own workers left out of the industry totals. The unit is the 94 minor groups of the Standard Occupational Classification, and every regression, over the supplements of 2013 to 2025, has group and year effects, is weighted by the group's 2012 employment, and clusters standard errors by group.

Almost all of the variation that identifies the estimate comes from one episode. Of the covariance between the instrument and employment growth, after the group and year effects, 70 per cent comes from the supplement of 2021 and 27 per cent from that of 2022. The Rotemberg weights of Goldsmith-Pinkham et al. (2020) place 0.45 on eating and drinking places and 0.15 on entertainment and recreation services, with beauty shops, hotels, personal services, and child day care next. With a single instrument the cluster-robust first-stage F is also the effective F of Olea and Pflueger (2013), and it is 14.0 on the full period, 6.4 without 2021, and 0.7 on the supplements of 2013 to 2019. I tried other alignments of growth with the supplement's reference periods (the March of the survey year against the previous calendar year, consecutive calendar years, and consecutive Marches), and none gives an F above 4 without 2020 to 2022. The estimate is therefore the response of occupational flows to the industry shock of 2020 and 2021, and in that episode the exclusion restriction is open to doubt. Unemployment

insurance that replaced more than the wage for many low earners, the risk to health, and the closure of schools and childcare all bore directly on departures out of employment in precisely the industries with the largest weight in the instrument.

4.3.2 Results

The instrument moves employment growth with a coefficient of 0.55 (standard error 0.15). Table 3 reports the second stage, with Anderson-Rubin 95 per cent sets, which remain valid if the instrument is weak. A log point of employment growth raised the change recorded in the account by 0.33 hundredths of last year's holders (0.06, Anderson-Rubin set 0.24 to 0.54). The account registers only about a third of the growth, a consequence of its windows. Last year's holders include everyone whose longest job of the previous calendar year was in the occupation, so a person hired during that year who is still in the job is a stayer, and the growth the account does not register consists largely of such hires.

Table 3: Response of each route to a log point of employment growth, per 100 holders last year, 94 occupational groups, supplements 2013 to 2025, instrumented with exposure to industry growth, standard errors clustered by group in parentheses

Route	Respondents			All records
	Response	Anderson-Rubin set	Share	Response
Change in the workforce over last year's holders	0.33 (0.06)	[0.24, 0.54]	100	0.32 (0.07)
Arrivals from another occupation	0.07 (0.03)	[0.02, 0.13]	20	0.04 (0.04)
Arrivals from education	0.01 (0.01)	[−0.01, 0.05]	3	0.01 (0.01)
Arrivals from other non-employment	0.04 (0.02)	[0.02, 0.08]	11	0.03 (0.02)
Departures to another occupation	−0.04 (0.02)	[−0.09, 0.01]	11	−0.03 (0.03)
Departures out of employment	−0.18 (0.04)	[−0.29, −0.12]	55	−0.20 (0.04)

Shares are each route's part of the response the account registers, in per cent, a fall in departures counting as a contribution. First stage 0.55 (0.15), effective $F = 14.0$, 1,222 group-years, 70 per cent of the identifying covariance from 2021 and 27 from 2022.

Of the response the account registers, more than half came through fewer departures out of employment (0.18, standard error 0.04), about a fifth through arrivals from other occupations (0.07, 0.03), a tenth through fewer departures to other occupations (0.04, 0.02), a tenth through arrivals from other non-employment (0.04, 0.02), and almost nothing through arrivals from education (0.01, 0.01). These are shares of the registered response, and the windows shape them. A departure out of employment requires only one week without work in March, while an arrival from outside employment requires a whole year without it, so the survey registers a fall in departures more readily than a rise in entry, and new entry during the previous year is among the hires that the account does not register at all. In the shock of 2020 and 2021, then, the occupations in the industries that contracted and recovered most lost and regained workers mainly at the margin of employment, the occupational counterpart of the finding of Blanchard and Katz (1992) that regions adjust at first through unemployment and participation.

The division is much the same in other specifications on the same episode. Leaving out 2014 and 2015, when the industry codes changed, gives a response of 0.30 (0.06) with 53 per cent through departures out of employment. At the detailed level, the 106 occupations with at least 150 respondents holding them in an average year give an effective F of 12.7 and a response

of 0.34 (0.07), with 58 per cent through departures out of employment and 14 through arrivals from another occupation. The 22 major groups give an effective F of 5.4 and a response of 0.36 (0.12). Without the leave-one-out correction, the instrument gives 0.18 (0.02) with the same division, although this is a variant of the same estimate and not an independent check, and least squares on employment growth gives 0.08 (0.02). Next year's instrument has no reduced-form relation with this year's change in the account (0.004, standard error 0.028, or 0.058 and 0.037 with this year's instrument held fixed). However, with this year's instrument held fixed it has a borderline relation with departures out of employment (-0.056 , standard error 0.028), so some of the measured fall in departures may reflect industries already on a different path.

Outside the episode the estimate has no content. Without 2021 the effective F is 6.4 and the response 0.40 (0.23), and without both 2020 and 2021 the Anderson-Rubin set for the response runs from 0.05 to 1.40, for departures out of employment from -0.58 to 0.07, and for arrivals from another occupation from 0.01 to 0.46, so nothing can be said about the division in ordinary years. The instrument's power also lies entirely in the lower-paid half of the groups, where the effective F is 17.1, and is nil in the upper half, so nothing can be learned from it about graduate occupations.

4.3.3 What the imputed records do to the estimate

The imputed records more than double the level of arrivals from other occupations, to 7.3 per hundred holders against 3.6 among respondents, without raising their response (Table 3). Relative to its level, the response of arrivals from another occupation is 1.85 per cent per log point of growth among respondents (standard error 0.69) and 0.57 (0.50) on all records, and the difference, estimated directly, is 1.28 (0.54), with an Anderson-Rubin set of 0.20 to 2.70. On all records, then, movement into an occupation appears both more frequent and less responsive to the shock. The difference rests on the same episode, however, as without 2020 and 2021 it is 2.23 (1.75), with a set from -1.3 to 8.8. For departures to another occupation the difference is -0.54 (0.40) and cannot be told from zero.

5 Limits

The prior occupation is the longest job of a calendar year and the current one is the job of a single week, so the count of people who held an occupation last year covers a twelve month window while the count of current workers is a moment. Thus the ratio between the two is not a growth rate, and in Section 4.3 the account registers about a third of an occupation's employment growth. Reconciling the flows with the change in an occupation's size requires a stock-flow treatment that takes the windows into account, and I do not attempt one here. Occupations also differ in seasonality and in how many jobs their people hold, and the longest job rule treats them differently in ways the account does not correct for.

As the survey looks back only one year, arrivals from outside employment include people returning to an occupation they had held before alongside those entering it for the first time.

Likewise, a departure is the statement that a person whose longest job last year was in the occupation is not in it in the survey week, which a single transition cannot separate from a temporary absence.

Movement between occupations is probably understated, for the reasons given in Section 3.3, and arrivals from education are understated by the definition examined in Section 4.2.2. The two errors work in opposite directions on the ratio of one to the other, and neither end of the range reported for it is a bound, as unreported moves would raise the ratio above 4.1 while entrants older than their early twenties would lower it below the wider definitions.

The origin of an arrival is where the person was, and the account does not record how anyone qualified. Training completed while in work appears in the lateral column, and in a licensed occupation a lateral arrival must already hold the licence, so the people in that column either held it already or earned it while working elsewhere. For this reason the account cannot rank training against other sources of an occupation's supply.

6 Conclusion

No published source gives how often people change occupation, or where each named occupation's workers come from and go, with a measure of precision, and the files that hold the answers contain a hazard. Records with an imputed supplement, a fifth of the recent files, generate about three quarters of the changes of occupation those files record, and the harmonised quality flag does not identify them. Anyone computing occupational mobility or transitions from the March supplement therefore requires the separate variable that identifies them.

The account I have built on respondents gives, for 214 occupations from 2009 to 2025, each recorded inflow and outflow with its standard error, and the full matrix of moves between occupations. Applied to the trend in mobility, it shows that the rise measured on all records after 2009 belongs mostly or wholly to the imputed records, whose share doubled, while respondents show no rise. Across occupations it agrees with the official transfer rates, which show no sign of the imputed records, although their levels cannot be checked from the published tables. For each occupation it also gives figures a planner can use: of a hundred registered nurses last year, six are no longer nurses in March, most of them not employed.

Applied to the industry shock of 2020 and 2021, the account shows the most exposed groups losing and regaining workers mainly through departures out of employment, a division the survey's reference periods help to shape and one that cannot be estimated outside that episode. Its two sides, however, cover different windows, so reconciling its flows with the change in an occupation's size requires a stock-flow treatment of those windows. And as two independent codings of a job recorded as unchanged disagree in about half of cases, the linked measure remains unusable at the level of the occupation pair.

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